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Abstract

We investigate the causes underlying the decline in the government expenditure multiplier after the Korean War. While this phenomenon has been documented before, we look at the decrease in relative multiplier values through the lens of a structural DSGE model, which we estimate using Bayesian methods and annual-frequency data from 1939 to 2017. The model replicates the observed fall in the expenditure multiplier; moreover, using a counterfactual exercise we show that the decline is accounted, for the most part, by changes in two of the model's structural parameters, namely, a decline in consumption habit persistence and a higher autocorrelation of the public expenditure processes. Taken together, these changes imply a stronger negative wealth effect (over consumption), a lower discretion of U.S. fiscal policy and, consequently, a multiplier of smaller magnitude.

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1 Introduction

In this paper, we investigate the causes underlying the decline in the government expenditure multiplier right after the Korean War. While this has been documented before (for example, see Hall 2009 or Barro and Redlick 2011), we look at the decrease in relative multiplier values through the lens of a structural dynamic stochastic general equilibrium (DSGE) model for the U.S. economy that we estimate using Bayesian methods.

Our model incorporates four features that the literature has shown to be relevant for analyzing fiscal multipliers. First, we differentiate between non-defense and defense expenditure; this distinction has been used in structural vector autoregression (SVAR) models to identify fiscal shocks (e.g., Hall 2009, Barro and Redlick 2011, and Ramey 2009, 2011), given that armed conflicts are a prime example of transitory public outlays. For this reason, we use an annual-frequency sample that runs from 1939 to 2017; of particular relevance is that we include World War II (hereafter, WWII) and the Korean War in our analysis—both of which involve massive fiscal expansions. As we explain below, we find it convenient to split the period of observations into two samples, one before and one after the Korean War.

Second, we consider a rich set of fiscal rules following Leeper, Plante, and Traum (2010). Under these rules, public expenditure and tax rates exhibit a certain degree of persistence and can respond both to business cycle conditions and to the state of debt. As public expenditure is disaggregated into defense and non-defense components, we take the former to be fully exogenous from output and debt; we require the latter to adjust when the defense budget is activated.[1]

Third, we incorporate anticipation of fiscal expenses into the model, following Schmitt-Grohé and Uribe (2012). We consider both "war news shocks" and non-defense news shocks, where the former represents anticipated changes in defense outlays and the latter is defined accordingly. While we share Ramey's (2011) strategy of using defense spending to identify the anticipated shocks,[2] our identification is based on a DSGE model, different from the narrative approach used in the literature.[3] Thus, the

[1] Hence, non-defense expenditure is more elastic than its defense counterpart, an assumption consistent with the conduction of fiscal policy during war times. In our model, when military expenses increase, the government budget automatically reduces non-defense spending.

[2] Ramey (2011) shows that studies employing SVARs generate biased estimates of the multiplier—unless a "defense news" variable is included in the estimation (see also Ramey and Zubairy 2018).

[3] In an illuminating example of the narrative approach, Ramey (2009) compiles "defense news" from *Business Week*

effects of anticipation remain orthogonal to other fiscal shocks. To our knowledge, we are the first to incorporate (and estimate) war news shocks in a DSGE framework.

Finally, our specification looks closely at the wealth effects that follow a fiscal expansion. For this purpose, we rely on two kinds of household preferences: one that follows King, Plosser, and Rebelo (1988, hereafter KPR preferences) with explicit wealth effects, and another based on Greenwood, Hercowitz, and Huffman (1988, hereafter GHH preferences) that eliminates these effects on the household's labor supply. In our environment, any increase in public expenditure generates a negative wealth effect that curbs leisure, so labor supply and output must increase. The parallel use of these two functional forms allows us to determine the importance of the wealth effect.

Altogether, our model successfully replicates the post-Korean War (hereafter, post-KW) fall in the expenditure multiplier. Using the policy functions from the model and our estimated parameters, we perform a counterfactual exercise, where we show that the decline in the magnitude of the multiplier is a consequence of a lower consumption habit persistence, paired with higher autocorrelation coefficients of the public expenditure processes. Taken together, these changes imply a stronger negative wealth effect (over consumption), a lower discretion of U.S. fiscal policy and, consequently, a multiplier of smaller magnitude.

Additional findings

We also document the following results. In Section 6.1, we show that our (DSGE-based) model does a good job in identifying the war news shocks, relative to the narrative defense news identified by Ramey (2011). The correlation coefficient between both series is positive: considering the exercise with GHH preferences, the correlation equals 0.48 for the pre-Korean War (hereafter, pre-KW) sample and 0.36 for the post-KW one. Our war news shocks correctly anticipate the expenditure increase of WWII, although the Korean War appears as a surprise. For the remaining war episodes, the shocks provided by Ramey (2011) and ours are fairly similar. Thus, we believe that our model is a valid framework to identify anticipated changes in defense (and non-defense) outlays in the U.S.

Even so, as shown in Section 6.4, we provide evidence that the role of anticipated fiscal shocks

and other newspapers across the 20th century, which allows her to assemble an estimate of the change in the expected present value of defense expenditure—a proxy for the beliefs about military build-ups. This variable is then integrated into an SVAR to build reliable estimates of the multiplier.

is ancillary at best. Through a variance decomposition exercise, we find that unanticipated military expenditure shocks generate substantial movements in output in the pre-KW sample, yet anticipated shocks are largely irrelevant. In the post-KW sample, most of the fluctuations in model variables can be accounted by total factor productivity (TFP) shocks; fiscal shocks (defense or non-defense, anticipated or not) play small roles in the variablity of model variables.

In Section 6.5, we show that distinguishing between defense and non-defense expenditure has a minor effect on the measurement of the fiscal multiplier: present-value multipliers (see Mountford and Uhlig 2009) based on defense spending are slightly smaller than those based on non-defense multipliers. This is a direct implication of the fiscal rules: changes in defense outlays trigger an internal budgetary mechanism in our model, so non-defense disbursements automatically adjust to smooth the impact of the fiscal expansion.

Lastly, we find that the use of alternative utility functions with (KPR) and without (GHH) wealth effects on household's labor supply do not affect the magnitude of the multiplier. If anything, multiplier estimates obtained under GHH preferences are slightly larger relative to a KPR specification.

Roadmap

The paper is structured as follows. Section 2 briefly surveys related research and Section 3 documents a first-pass concerning the decline in the government multiplier. We present our model in Section 4, while Section 5 explains our parametrization strategy and reports the outcome of our estimation. Section 6 lays out our results: the identified war news shocks, impulse-response functions, variance decompositions, and our measured fiscal multipliers. Section 7 contains several counterfactual exercises that help understand why the multiplier has decreased after the Korean War. Section 8 wraps up with our concluding remarks.

2 Connection with the Literature

The fiscal multiplier is defined as the change in output relative to a discretionary fiscal variation; i.e., a change in non-automatic fiscal incentives.[4] However, to this day, the literature has not reached a

[4] The original definition of the multiplier dates back to Keynes's ideas—later developed by the Keynesian school—through the concept of marginal propensity to consume (MPC). As usually explained in introductory textbooks, since MPC $\in (0, 1)$ the multiplier necessarily exceeds unity, given that $1/(1 - \text{MPC}) > 1$.

consensus regarding the empirical value of the fiscal multiplier.

To see this, consider the work of Ilzetzki, Mendoza, and Végh (2013), who estimate spending multipliers for a set of 44 countries using an SVAR panel, taking note of how the values change along different economic dimensions. First, they estimate impact multipliers that range from -0.03 (not statistically significant) for developing countries to 0.39 for high-income countries; for the latter group, the multipliers can reach 0.66 in the long run. Second, they find that the multiplier varies from 0.15 on impact to 1.4 in the long run for countries under de-facto fixed exchange rate regimes; under flexible regimes, the multipliers are negative at all time horizons (also not statistically significant). Third, they show that multipliers are inversely related to the volume of foreign trade: impact multipliers range between 0.6 and 1.1 for relatively closed countries. Finally, they provide evidence that the impact multiplier is nearly zero for highly indebted countries, but can become very negative (as low as -3) in the long run: when debt levels are high, increases in government expenditure signal that fiscal tightening will be required in the near future.

In light of these results, it is not clear whether the circumstances that generate high multipliers are present in economies such as the U.S.; nonetheless, many authors *have* estimated the U.S. expenditure multiplier using different toolkits, of which we discuss three: classical econometrics (OLS), time-series econometrics (SVARs), and DSGE models.

In the first category, we can include Barro and Redlick (2011) and Hall (2009). In particular, Hall concludes that the multiplier may range from 0.7 to unity for the U.S. economy; he argues that a multiplier value of 1.7 could be reached if the nominal interest rate hits the zero lower bound.

SVARs have been used extensively in recent years; among many others, we list the contributions of Blanchard and Perotti (2002); Galí, López-Salido, and Vallés (2007); Perotti (2008); Mountford and Uhlig (2009); Ramey (2011); Auerbach and Gorodnichenko (2012); and Ramey and Zubairy (2018). Aside from the last three papers in this list, these multiplier estimates are based on orthogonalization conditions that generate a set of structural shocks from the SVAR forecasting errors.[5]

Several of these SVAR-based studies use defense expenditure as a tool to identify fiscal shocks, given that it is considered exogenous to output. Under this identification scheme, the impulse-response functions from fiscal shocks are used to estimate the value of the multiplier, and yield peak output

[5] Most—but not all—of these estimates are what Ramey (2016, p. 116) calls the "peak output response to the intial govenment spending impact effect," as opposed to the present value approach favored by Mountford and Uhlig (2009).

responses that range between 0.3 and 0.9 on impact and reach values above unity after several quarters.

However, Ramey (2011) warns that a portion of fiscal shocks identified using SVAR tools might be anticipated, so that they cannot be used to assess the fiscal multiplier. Instead, she uses the defense news collected from her narrative approach: as long as the series reflects beliefs about future changes in military spending, they allow her to identify unanticipated fiscal shocks. She estimates peak output response multipliers between 1.1 and 1.2.[6] Auerbach and Gorodnichenko (2012) rely on the identification strategy of Blanchard and Perotti (2002) and the defense news series of Ramey (2011) to calculate multipliers during expansions and recessions; they find that using Ramey's defense news time series generates larger values in both cases. Ramey and Zubairy (2018) elaborate on the work of Ramey (2011) by first building a quarterly-frequency GDP series that runs from 1889 to 2015. Using this expanded dataset, they ask whether the size of the multiplier depends on slack conditions (e.g., high vs. low unemployment) or the presence of a zero lower bound. In general, they cannot find evidence for large multipliers in the former scenario, but do find multipliers above unity when certain conditions are met in the latter one.

Finally, researchers have also used DSGE models to calculate expenditure multipliers. For instance, Leeper et al. (2010) conclude that the estimated multiplier can vary widely, depending on the sensitivity of the fiscal instrument with respect to debt. Cogan, Cwik, Taylor, and Wieland (2010) show that standard NK models (e.g., Smets and Wouters 2007) cannot generate multipliers above unity with a permanent increase in government expenditure. They ask whether adding non-Ricardian consumers (Galí et al. 2007) can produce higher values, but find it cannot. Zubairy (2014) uses a model with nominal frictions, deep habits (see Ravn, Schmitt-Grohé, and Uribe 2006) and a rich fiscal policy block with automatic stabilizers to find multipliers that are above unity in the short run. (Neither of these DSGE models considers anticipation effects.)

[6] See also the work of Ben Zeev and Pappa (2015). They find that unexpected increases in military expenditure push TFP and output upwards, generating a multiplier of 0.94, but that this response goes to zero once the TFP channel is shut down.

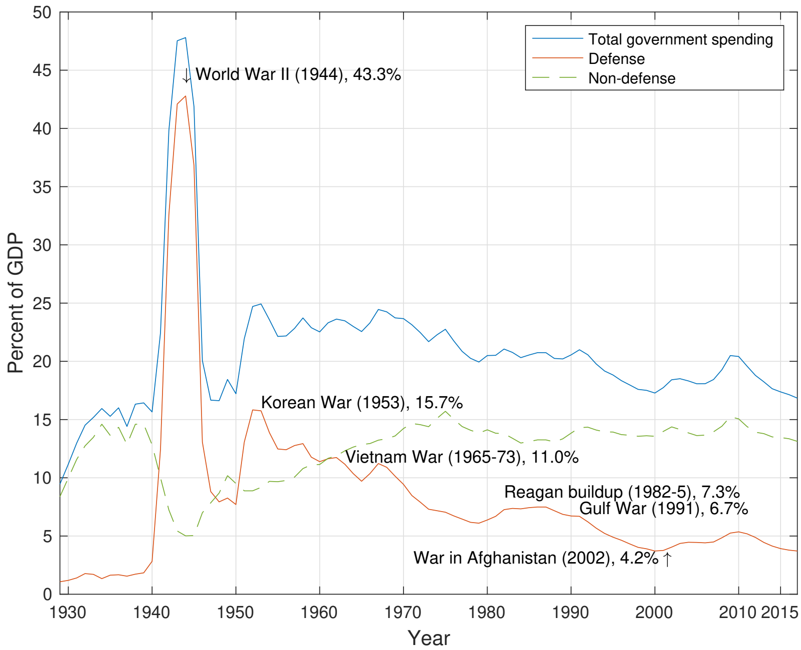

Figure 1: Government expenditure (total, defense, and non-defense), 1929–2017.

3 Evidence of a Declining Multiplier

Figure 1 presents defense, non-defense, and total government expenditure for the U.S., from 1929 to 2017.[7] We borrow two lessons from these series. First, the years surrounding WWII (1941–1945) exhibit an unprecedented increase in military spending. A few years later, the Korean War also brings an increase in this variable, but of smaller magnitude relative to WWII (defense expenditure represents a bit over 43% of GDP in 1944; its 1953 value is considerably lower, at 16% of GDP). Nonetheless, defense expenditure accounts for the bulk of the transitory component in total government expenditure. Second, there is a budgetary tradeoff involving defense and non-defense expenditure—particularly evident in the decades after the Korean War.

[7] As detailed in Appendix A.1, our defense expenditure series considers federal-level data, while non-defense expenditure considers both the federal and state and local levels. Total government expenditure is the sum of defense and non-defense, as defined above.

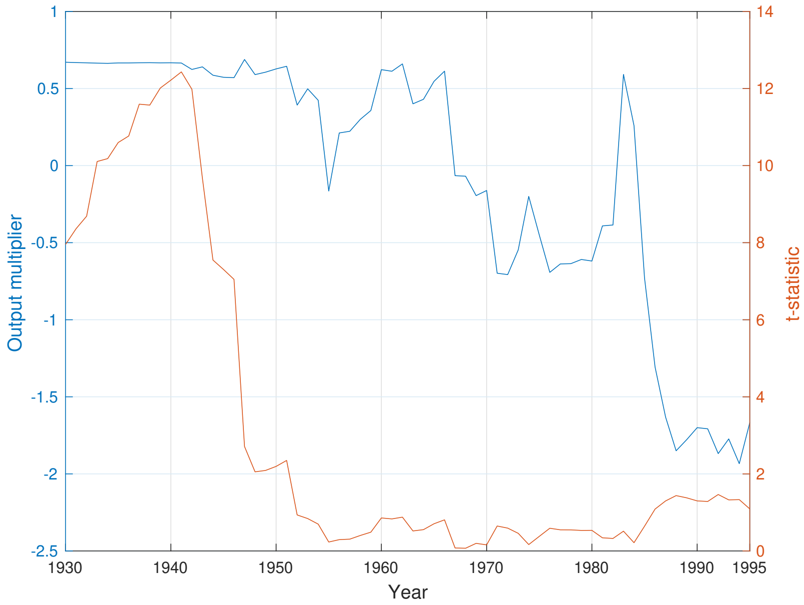

Figure 2: Output multiplier and associated t-statistic.

A first pass at the expenditure multiplier A straightforward way to quantify the relation between output and defense expenditure is the regression

$$\frac{\Delta Y_t}{Y_{t-1}} = a + m_Y \frac{\Delta M_t}{Y_{t-1}} + e_t, \tag{3.1}$$

where Y_t and M_t denote real GDP and defense expenditure at time t. This specification is used by Hall (2009) and Barro and Redlick (2011), who show that the OLS estimator m_Y can be interpreted as a weighted-value output multiplier

$$m_Y = \frac{\operatorname{cov}(\Delta Y_t/Y_{t-1}, \Delta M_t/Y_{t-1})}{\operatorname{var}(\Delta M_t/Y_{t-1})} = \frac{\sum_t \left(\frac{\Delta M_t}{Y_{t-1}}\right)\left(\frac{\Delta Y_t}{Y_{t-1}}\right)}{\sum_t \left(\frac{\Delta M_t}{Y_{t-1}}\right)^2} = \sum_t \omega_t \frac{\Delta Y_t}{\Delta M_t},$$

with weight $\omega_t \equiv \left(\frac{\Delta M_t}{Y_{t-1}}\right)^2 / \sum_t \left(\frac{\Delta M_t}{Y_{t-1}}\right)^2$.

We rework Hall's exercise by estimating a sequence of regressions in a moving window that begins in 1930 and changes the starting year all the way to 1995. (In all regressions, the sample ends in 2017.) Figure 2 presents the estimates of the output multiplier m_Y, 66 multipliers for 66 shrinking samples, along with their corresponding t-statistic.[8,9] Overall (and consistent with Hall 2009 and Barro and Redlick 2011), we find fiscal multipliers around 0.7, but only when our sample includes WWII and the Korean War. In our second sample, the fiscal multiplier estimates are not statistically significant—so they must be taken as zero.

Guided by this evidence, we split the sample in two: a first sample that considers 1939 to 1954—which includes both WWII and the Korean War—and another from 1955 to 2017, a period with war episodes whose volume of spending are of minor importance compared to those of the first one.

4 A Model to Measure the Expenditure Multiplier

Our model economy consists of three kinds of agents: a representative household, a representative firm, and a government.

4.1 Household

The representative household's preferences depends on (current and past) consumption and hours worked. Household decisions are constrained by a budget: gross income must be distributed between consumption, saving, and taxes. Saving can be done through a government bond or physical capital. Our model also includes habit formation in consumption and capital adjustment costs. In particular, the household chooses sequences of consumption c_t, labor supply n_t, investment x_t, utilization rate of capital u_t, and bond holdings b_t to solve the following problem:

$$\max \quad E_0 \sum_{t=0}^{\infty} \beta^t U(c_t, c_{t-1}, n_t) \tag{4.1}$$

$$\text{s.t.} \quad c_t + x_t + b_t = r_t u_t k_t + w_t n_t + r^b_{t-1} b_{t-1} - \zeta_t \tag{4.2}$$

$$\zeta_t = \tau_{kt} r_t u_t k_t + \tau_{nt} w_t n_t + \tau_{ct} c_t - \tau_t - \delta(u_t)\tau_{kt} k_t \tag{4.3}$$

$$k_{t+1} = [1 - \delta(u_t)]k_t + x_t[1 - S(x_t/x_{t-1})]. \tag{4.4}$$

[8] The first multiplier (1930–2017) is estimated as 0.7, while the last estimate is -1.6, corresponding to 1995–2017.

[9] As a robustness check, we repeat this experiment but change the direction of the rolling window: for each regression, one final year of the sample is added, taking the starting window of 1930 to 1950 as fixed. In this case, the multiplier hovers around 0.67 while the statistical significance of these estimates displays an upward trend (as one year is added at the end of the sample) that is always above the critical threshold.

In the above, $\beta \in (0,1)$ is a discount factor and the utility function U is strictly increasing, strictly concave, and twice continuously differentiable. Equation (4.2) represents the household's budget constraint: the stock of physical capital is denoted by k_t, which is rented to the representative firm at a rental rate r_t. The real wage is denoted by w_t. In addition, r_t^b denotes the gross rate of return on government bonds (the rate is fixed in period $t-1$) and ζ_t denotes the fiscal policy component of the household budget. Following (4.3), the nonnegative processes $\{\tau_{kt}, \tau_{nt}, \tau_{ct}\}_{t=0}^{\infty}$ represent tax rates levied over capital income, labor income, and the purchase of consumption goods, respectively, while τ_t denotes government transfers to the household. The last term in (4.3) represents a depreciation allowance. Equation (4.4) is the law of motion for capital, where the depreciation function δ depends on capital utilization and there are investment adjustment costs parametrized by the function S, which satisfies $S(1) = S'(1) = 0$ and $S''(1) > 0$.

As usual, we assume that the household behaves competitively and takes the processes for prices $\{r_t, w_t, r_t^b\}_{t=0}^{\infty}$ and fiscal policy $\{\tau_{kt}, \tau_{nt}, \tau_{ct}, \tau_t\}_{t=0}^{\infty}$ as given.

Functional forms: preferences The utility function (4.1) can take on two functional forms. First, an expression that is consistent with KPR preferences:

$$U(c_t, c_{t-1}, n_t) = \frac{(c_t - \mu c_{t-1})^{1-\gamma}}{1-\gamma} - \phi \frac{n_t^{1+\chi}}{1+\chi}, \tag{4.5}$$

where $\gamma > 0$ is the coefficient of relative risk aversion, $\mu \in (0,1)$ determines the strength of the consumption habit persistence, $\phi > 0$ is a labor disutility weight, and $\chi > 0$ denotes the (inverse of the) Frisch elasticity of labor supply. The second functional form relies on GHH preferences instead:[10]

$$U(c_t, c_{t-1}, n_t) = \frac{1}{1-\gamma}\left(c_t - \mu c_{t-1} - \phi \frac{n_t^{1+\chi}}{1+\chi}\right)^{1-\gamma}. \tag{4.6}$$

[10] We decide to use GHH preferences—as opposed to more general functional forms, for example those in Jaimovich and Rebelo (2009)—based on the empirical findings of Schmitt-Grohé and Uribe (2012, Table II, p. 2750). Their estimation results strongly suggest that a GHH functional form better fits postwar U.S. data. (See the estimates for parameter γ; starting from a uniform prior of 0.5, the posterior is revised to a value of 0.00, consistent with GHH preferences.)

Functional forms: depreciation and adjustment costs We impose the following functional forms for the depreciation and investment adjustment cost functions:

$$\delta(u_t) = \delta_0 + \delta_1(u_t - 1) + \frac{\delta_2}{2}(u_t - 1)^2 \tag{4.7}$$

$$S\left(\frac{x_t}{x_{t-1}}\right) = \frac{\kappa}{2}\left(\frac{x_t}{x_{t-1}} - 1\right)^2, \tag{4.8}$$

where $\delta_0, \delta_1, \delta_2, \kappa > 0$. These functional forms are used frequently in the business cycle literature (e.g., Schmitt-Grohé and Uribe 2012).

4.2 Firm

The representative firm rents capital K_t and labor services N_t from the household and transforms them into output y_t. We assume that the firm behaves competitively, takes the processes for the rental prices $\{r_t, w_t\}_{t=0}^{\infty}$ as given, and solves

$$\begin{aligned} \max \quad & y_t - r_t K_t - w_t N_t \\ \text{s.t.} \quad & y_t = z_t K_t^{\alpha} N_t^{1-\alpha}. \end{aligned}$$

In the above, $\alpha \in (0, 1)$ represents the capital share and z_t is a stationary TFP disturbance. We assume that z_t follows

$$\log(z_t) = \rho_z \log(z_{t-1}) + \varepsilon_{zt}, \tag{4.9}$$

where $\rho_z \in (0, 1)$ is a persistence parameter and ε_{zt} is an i.i.d. process with zero mean and standard deviation σ_z.

4.3 Government and fiscal rules

The government trades bonds B_t with the household and levies taxes over consumption and capital and labor incomes. It uses these resources to finance bond payments $r_{t-1}^{b} B_{t-1}$ and exogenous sequences of military purchases m_t, non-military consumption g_t, transfers τ_t, and depreciation allowances $\delta(u_t)\tau_{kt}k_t$.

The government's budget constraint is

$$g_t + m_t + r^b_{t-1}B_{t-1} + \tau_t + \delta(u_t)\tau_{kt}k_t = B_t + \tau_{ct}c_t + \tau_{kt}r_t u_t k_t + \tau_{nt}w_t n_t. \tag{4.10}$$

We specify a set of fiscal rules for public expenditure, transfers, and tax rates. Consistent with Figure 1, we assume that defense outlays do not respond to output nor to the state of public debt. Rather, defense expenses are driven by global geopolitical factors and not by domestic economic conditions; hence, we posit the law of motion

$$\widehat{m}_t = \rho_m \widehat{m}_{t-1} + \varepsilon_{mt} + \varepsilon^{\text{war}}_{m,t-1}, \tag{4.11}$$

where $\widehat{m}_t$ denotes the percent deviation of military expenditure from its steady state value. (This notation will be used extensively in what follows.) Equation (4.11) indicates that defense expenses evolve with relative persistency ρ_m and that deviations from steady state can be attributed to a non-anticipated fiscal shock ε_{mt} or the war news shock $\varepsilon^{\text{war}}_{m,t-1}$, which becomes known one period (in our case, a year) in advance. The war news shock is key to properly measuring the effects of military expenditure on the economy, as changes in the variable are often anticipated. Both innovations are i.i.d. processes with zero mean and standard deviation σ_m and σ_{war}, respectively.

Non-defense expenditure follows an augmented version of the fiscal rule suggested by Leeper et al. (2010):

$$\widehat{g}_t = \rho_g \widehat{g}_{t-1} - \theta^m_g \widehat{m}_t - \theta^y_g \widehat{y}_{t-1} - \theta^b_g \widehat{b}_{t-1} + \varepsilon_{gt} + \varepsilon^{\text{gov}}_{g,t-1}. \tag{4.12}$$

From (4.12), $\widehat{g}_t$ adjusts automatically to past values and in response to defense outlays, the business cycle, and the state of debt. Hence, there is a systematic correction in non-defense disbursements when either defense spending differs from its steady-state value, output fluctuates around its steady state position, or debt accumulates relative to the stationary state. The response parameters $\{\theta^m_g, \theta^y_g, \theta^b_g\}$ are expected to be positive. Finally, non-defense expenditure is driven by a fiscal shock ε_{gt} and an anticipated government spending shock $\varepsilon^{\text{gov}}_{g,t-1}$—also known one period in advance—which captures the expected changes that will occur to government expenditure in the coming period. Both innovations are i.i.d. processes with zero mean and standard deviations σ_g and σ_{gov}, respectively.

The remaining fiscal rules are specified as in Leeper et al.:

$$\widehat{\tau}_t = \rho_\tau \widehat{\tau}_{t-1} - \theta_\tau^y \widehat{y}_{t-1} - \theta_\tau^b \widehat{b}_{t-1} + \varepsilon_{\tau t}, \tag{4.13}$$

$$\widehat{\tau}_{kt} = \rho_k \widehat{\tau}_{k,t-1} + \theta_k^y \widehat{y}_{t-1} + \theta_k^b \widehat{b}_{t-1} + \varepsilon_{kt}, \tag{4.14}$$

$$\widehat{\tau}_{nt} = \rho_n \widehat{\tau}_{n,t-1} + \theta_n^y \widehat{y}_{t-1} + \theta_n^b \widehat{b}_{t-1} + \varepsilon_{nt}, \tag{4.15}$$

$$\widehat{\tau}_{ct} = \rho_c \widehat{\tau}_{c,t-1} + \varepsilon_{ct}. \tag{4.16}$$

In the above, $\{\varepsilon_{\tau t}, \varepsilon_{kt}, \varepsilon_{nt}, \varepsilon_{ct}\}$ represent i.i.d. non-anticipated fiscal shocks with zero mean and standard deviations σ_j, $j = \tau, k, n, c$. Unlike Leeper et al., we assume that the non-anticipated fiscal shocks are all orthogonal. Except for the consumption tax rate, there is an automatic response with respect to output fluctuations and to the level of debt: the response parameters $\left\{\theta_\tau^y, \theta_\tau^b, \theta_k^y, \theta_k^b, \theta_n^y, \theta_n^b\right\}$ are expected to take on positive values.

4.4 Aggregate feasibility

Finally, the feasibility constraint of the economy dictates that output must be either consumed (by households or by the government) or invested:

$$y_t = c_t + x_t + g_t + m_t. \tag{4.17}$$

Appendix A.2 characterizes the equilibrium conditions of the model.

5 Parametrization and Estimation

In this section, we discuss the parametrization of our model. We first calibrate a subset of parameters and steady state values using sample averages, and then use Bayesian techniques to estimate the remaining ones. Appendix A.3 presents additional details regarding the estimation procedure.

5.1 Calibrated parameters

Table 1 displays parameters determined ex-ante and target values that the model aims to reproduce in the steady state. The top panel shows calibrated parameters. The capital income share is set to $\alpha = 1/3$

Table 1: Parameters determined ex-ante and target values.

Parameter or target ratio	1939–54	1955–17
Capital income share, α	0.333	0.333
Capital income tax rate, τ_k	0.344	0.246
Labor income tax rate, τ_n	0.108	0.191
Consumption tax rate, τ_c	0.051	0.052
Consumption tax persistency, ρ_c	0.310	0.724
Consumption tax volatility, σ_c	0.071	0.027
Consumption-to-output	0.542	0.544
Investment-to-output	0.199	0.250
Non-defense expenditure-to-output	0.088	0.135
Defense expenditure-to-output	0.172	0.071
Transfers-to-output	0.031	0.095
Debt-to-output	0.771	0.549
Market-to-total hours	0.300	0.300

for both samples, while the stationary tax rates are calculated as an average of calculated values (see Appendix A.1 for details on data construction).[11] The bottom panel shows our targets for the model economy. Between samples, the main difference concerns the composition and magnitude of public expenditure. Consumption relative to output is stable around 0.54, while investment moves opposite to government spending.

5.2 Bayesian estimation

We estimate the remaining parameters using Bayesian techniques. These series are assumed observable: consumption c_t; investment x_t; defense and non-defense expenditure $\{m_t, g_t\}$; transfers τ_t; tax rates on capital income, labor income, and consumption $\{\tau_{kt}, \tau_{nt}, \tau_{ct}\}$; and debt b_t.[12]

5.2.1 Household and firm parameters

Tables 2a (considering KPR preferences) and 2b (GHH preferences) present the prior and posterior means of household and firm parameters. Overall, parameter estimates seem robust to the different preference specifications, though a subset of these change as we move from pre- to post-KW samples.

For the KPR case, two parameters remain practically constant between samples: the quadratic

[11] The parameters associated with the consumption tax (ρ_c and σ_c) are computed with a simple OLS regression.

[12] We add measurement errors to defense and non-defense expenditure, transfers, consumption, and investment. See Appendix A.1 for details on the observable variables used in the estimation procedure.

Table 2a: Prior and posterior distributions for model parameters (household and firm), KPR preferences.

Parameter	Prior distribution			Posterior distribution for sample						Relative
				1939–54			1955–17			
	Density	Mean	SD	Mean	95% interval		Mean	95% interval		
Consumption habit persistence, μ	Beta	0.50	0.20	0.856	0.717	0.979	0.536	0.331	0.730	1.6
Inverse Frisch elasticity, χ	Gamma	2.00	0.50	2.224	1.274	3.235	1.821	0.928	2.753	1.2
Intertemporal elasticity of substitution, γ	Gamma	1.75	0.50	1.980	0.999	3.044	1.569	0.774	2.422	1.3
Depreciation rate, constant term, δ_0	Beta	0.05	0.03	0.032	0.002	0.069	0.068	0.008	0.135	0.5
Depreciation rate, quadratic term, δ_2	Gamma	1.00	0.30	1.287	0.697	1.965	1.288	0.744	1.887	1.0
Investment adjustment cost, κ	Gamma	5.00	0.25	4.973	4.493	5.467	4.978	4.499	5.460	1.0

Table 2b: Prior and posterior distributions for model parameters (household and firm), GHH preferences

Parameter	Prior distribution			Posterior distribution for sample						Relative
				1939–54			1955–17			
	Density	Mean	SD	Mean	95% interval		Mean	95% interval		
Consumption habit persistence, μ	Beta	0.50	0.20	0.738	0.494	0.933	0.406	0.216	0.597	1.8
Inverse Frisch elasticity, χ	Gamma	2.00	0.50	2.238	1.363	3.172	1.856	1.005	2.782	1.2
Intertemporal elasticity of substitution, γ	Gamma	1.75	0.50	1.613	0.852	2.449	1.573	0.784	2.435	1.0
Depreciation rate, constant term, δ_0	Beta	0.05	0.03	0.029	0.002	0.060	0.063	0.007	0.130	0.5
Depreciation rate, quadratic term, δ_2	Gamma	1.00	0.30	1.339	0.762	1.945	1.183	0.708	1.696	1.1
Investment adjustment cost, κ	Gamma	5.00	0.25	4.983	4.507	5.473	4.959	4.480	5.455	1.0

depreciation term δ_2 (with posterior mean of 1.29) and the adjustment cost parameter κ (with estimates of 4.97 and 4.98). The steady-state depreciation parameter δ_0 increases from 0.03 to 0.07, while the consumption habit persistence parameter μ and the relative risk aversion coefficient γ fall from 0.86 to 0.54 and from 1.98 to 1.57, respectively. The Frisch elasticity of labor supply (given by $1/\chi$) rises from 0.45 to 0.55 between samples.[13] The GHH estimation results tell a somewhat similar story: the adjustment cost parameter κ remains constant between samples yet other parameters change. The habit persistence (μ; 0.74 to 0.41), constant relative risk aversion (γ; 1.61 to 1.57), and quadratic depreciation (δ_2; 1.34 to 1.18) parameters fall, while the Frisch elasticity ($1/\chi$; 0.45 to 0.54) and steady-state depreciation (δ_0; 0.03 to 0.06) parameters increase.

Leeper et al. (2010) document similar values for these parameters, noting that they assume KPR preferences and use quarterly data and a different time sample (1976–2008). They estimate a posterior mode for consumption habit persistence between 0.5 and 0.7; this range of values coincides with the estimates found in Christiano, Eichenbaum, and Evans (2005), who use a DSGE model with nominal rigidities. Our estimate is also in range with those provided by Havranek, Rusnak, and Sokolova (2017), who survey 81 research papers that have estimated the consumption habit persistence parameter for different economies and applying different methodologies. As we show later, the expenditure multipliers are highly sensitive to changes in this parameter: the higher the value of μ, the lower the household response to fiscal policies and the lower the wealth effect on consumption. Thus, lower values of this parameter reduce the output response to government spending.[14]

Regarding the intertemporal elasticity of substitution γ and the inverse Frisch elasticity χ, Leeper et al. estimate a posterior mode within $[2.3, 2.5]$ and $[1.8, 2.1]$, respectively. Our posterior mean estimate for γ is rather low at 1.57, though our value for χ is in line with their estimates.

We combine the targets reported in Table 1 and our estimation results to calibrate additional parameters for the two samples. Table 3 summarizes our calcuations. With exception of the labor disutility weight parameter ϕ, the results reported in the tables are similar regardless of the utility function used. The (gross) long-run bond treasury rate r^b hovers around 1.02 for both samples, which implies subjective

[13] Heathcote, Storesletten, and Violante (2010) and Chetty, Guren, Manoli, and Weber (2011)—who use a married couple as the notion of household—propose an estimate of 0.72.

[14] The second-sample estimate $\mu = 0.54$ is similar to the values reported by other authors (see Havranek et al. 2017 for a survey), even under alternative model specifications.

Table 3: Parameters that require solving the model and implied steady state values.

Parameter or steady state value	KPR preferences		GHH preferences	
	1939–54	1955–17	1939–54	1955–17
Subjective discount rate, β	0.986	0.983	0.988	0.984
Linear term, depreciation rate, δ_1	0.054	0.090	0.048	0.084
Labor disutility weight, ϕ	891.9	85.83	6.005	5.727
Capital-output ratio, K/Y	6.190	3.696	6.972	3.965
Treasury bond gross rate, r^b	1.014	1.017	1.013	1.016
Rental price of capital, r	0.054	0.090	0.048	0.084

discount rates β in the neighborhood of 0.98. The capital-output ratio implied by KPR preferences is lower than the one associated with GHH utility (6.19 vs. 6.97) for the pre-KW sample, but fairly similar for the second one (3.70 vs. 3.97). Though they have different magnitudes—a consequence of the utility form—the labor disutility weight falls between periods.

5.2.2 Fiscal policy parameters

Tables 4a (KPR) and 4b (GHH) report the prior and posterior means of the remaining model parameters; we believe three results are worth noting.

First, the persistency of fiscal rules and TFP shocks—as deduced from the autocorrelation coefficients—has increased following the end of the Korean War. These estimates meet those reported by Leeper et al. (2010); using quarterly data and making no distinction between defense and non-defense expenditure, they document $\rho_{g+m} = 0.97$.[15] Our estimate is also consistent with that of Afonso, Agnello, and Furceri (2010), who estimate a persistence of 0.83 using annual data for 1980–2007. As we illustrate later, the rise in the autocorrelation coefficients of public expenditure processes also accounts for the fall in the multiplier.

Second, the response parameters (with the exception of θ_k^b and θ_k^b) have decreased from the pre- to the post-KW sample. Parameter θ_g^m—the response of non-defense expenditure to military outlays—falls from 0.11 to 0.08: thus, post-KW, a one-dollar increase in military spending induces an automatic adjustment of 8 cents in non-defense outlays. The responsiveness of non-defense expenditure with respect to output is estimated as $\theta_g^y = 0.07$ in the post-KW sample; this widely contrasts with estimates in other

[15] Note that the annual frequency of our dataset implies $\rho_g^{1/4} = 0.826^{1/4} = 0.953$; similarly, $\rho_m^{1/4} = 0.973$.

Table 4a: Prior and posterior distributions for model parameters (exogenous processes), KPR preferences.

Parameter	Prior distribution			Posterior distribution for sample						Relative
				1939–54			1955–17			
	Density	Mean	SD	Mean	95% interval		Mean	95% interval		
AR coefficient, ρ_z	Beta	0.70	0.20	0.340	0.087	0.607	0.554	0.382	0.718	0.6
AR coefficient, ρ_m	Beta	[a]	[b]	0.456	0.412	0.500	0.896	0.798	0.985	0.5
AR coefficient, ρ_g	Beta	[a]	[b]	0.423	0.382	0.465	0.826	0.673	1.000	0.5
AR coefficient, ρ_τ	Beta	[a]	[b]	0.465	0.420	0.511	0.808	0.723	0.889	0.6
AR coefficient, ρ_k	Beta	[a]	[b]	0.487	0.440	0.535	0.742	0.673	0.810	0.7
AR coefficient, ρ_n	Beta	[a]	[b]	0.306	0.277	0.336	0.624	0.565	0.682	0.5
Response, θ_g^m	Gamma	0.10	0.03	0.108	0.056	0.164	0.086	0.041	0.133	1.3
Response, θ_g^y	Gamma	0.07	0.05	0.072	0.002	0.171	0.065	0.001	0.156	1.1
Response, θ_τ^y	Gamma	0.20	0.10	0.206	0.037	0.407	0.164	0.032	0.327	1.3
Response, θ_k^y	Gamma	1.00	0.30	0.797	0.374	1.263	0.903	0.448	1.392	0.9
Response, θ_n^y	Gamma	0.50	0.25	0.671	0.177	1.218	0.372	0.082	0.705	1.8
Response, θ_g^b	Gamma	0.40	0.20	0.250	0.109	0.400	0.091	0.026	0.164	2.8
Response, θ_τ^b	Gamma	0.40	0.20	0.144	0.036	0.260	0.123	0.049	0.202	1.2
Response, θ_k^b	Gamma	0.40	0.20	0.181	0.042	0.338	0.318	0.176	0.468	0.6
Response, θ_n^b	Gamma	0.40	0.20	0.287	0.091	0.493	0.123	0.053	0.195	2.3
SD, σ_z	Inverse Gamma	0.05	0.02	0.057	0.026	0.095	0.034	0.025	0.044	1.7
SD, σ_m	Inverse Gamma	[c]	[d]	0.470	0.426	0.515	0.046	0.042	0.051	10.1
SD, σ_{war}	Gamma	0.05	0.02	0.051	0.016	0.093	0.034	0.013	0.055	1.5
SD, σ_g	Inverse Gamma	[c]	[d]	0.055	0.050	0.060	0.020	0.018	0.022	2.7
SD, σ_{gov}	Gamma	0.05	0.02	0.036	0.013	0.062	0.027	0.013	0.041	1.3
SD, σ_τ	Inverse Gamma	[c]	[d]	0.196	0.178	0.215	0.036	0.033	0.039	5.4
SD, σ_k	Inverse Gamma	[c]	[d]	0.187	0.169	0.205	0.066	0.061	0.072	2.8
SD, σ_n	Inverse Gamma	[c]	[d]	0.158	0.143	0.173	0.035	0.032	0.038	4.5

Notes. [a] Uses the OLS estimate of the corresponding coefficient in the policy rule. [b] Uses 0.05×OLS estimate of the corresponding coefficient in the policy rule. [c] Uses the standard deviation of the residual from the estimated policy rule. [d] Uses 0.05×SD of the residual from the estimated policy rule.

Table 4b: Prior and posterior distributions for model parameters (exogenous processes), GHH preferences.

Parameter	Prior distribution			Posterior distribution for sample 1939–54			Posterior distribution for sample 1955–17			Relative
	Density	Mean	SD	Mean	95% interval		Mean	95% interval		
AR coefficient, ρ_z	Beta	0.70	0.20	0.355	0.096	0.623	0.613	0.432	0.788	0.6
AR coefficient, ρ_m	Beta	[a]	[b]	0.458	0.414	0.502	0.899	0.803	0.987	0.5
AR coefficient, ρ_g	Beta	[a]	[b]	0.423	0.382	0.465	0.827	0.678	1.000	0.5
AR coefficient, ρ_τ	Beta	[a]	[b]	0.466	0.420	0.512	0.808	0.721	0.889	0.6
AR coefficient, ρ_k	Beta	[a]	[b]	0.487	0.440	0.534	0.740	0.672	0.809	0.7
AR coefficient, ρ_n	Beta	[a]	[b]	0.306	0.276	0.336	0.623	0.566	0.681	0.5
Response, θ_g^m	Gamma	0.10	0.03	0.108	0.057	0.164	0.084	0.041	0.132	1.3
Response, θ_g^y	Gamma	0.07	0.05	0.070	0.002	0.166	0.065	0.002	0.154	1.1
Response, θ_τ^y	Gamma	0.20	0.10	0.204	0.036	0.402	0.160	0.031	0.312	1.3
Response, θ_k^y	Gamma	1.00	0.30	0.774	0.355	1.220	0.912	0.446	1.397	0.8
Response, θ_n^y	Gamma	0.50	0.25	0.689	0.192	1.228	0.379	0.094	0.703	1.8
Response, θ_g^b	Gamma	0.40	0.20	0.255	0.119	0.407	0.087	0.027	0.154	2.9
Response, θ_τ^b	Gamma	0.40	0.20	0.148	0.040	0.268	0.127	0.051	0.208	1.2
Response, θ_k^b	Gamma	0.40	0.20	0.187	0.047	0.341	0.320	0.178	0.467	0.6
Response, θ_n^b	Gamma	0.40	0.20	0.284	0.092	0.487	0.121	0.051	0.194	2.4
SD, σ_z	Inverse Gamma	0.05	0.02	0.052	0.027	0.084	0.030	0.022	0.038	1.8
SD, σ_m	Inverse Gamma	[c]	[d]	0.470	0.425	0.515	0.046	0.042	0.051	10.1
SD, σ_{war}	Gamma	0.05	0.02	0.052	0.015	0.094	0.034	0.013	0.055	1.6
SD, σ_g	Inverse Gamma	[c]	[d]	0.055	0.050	0.060	0.020	0.018	0.022	2.7
SD, σ_{gov}	Gamma	0.05	0.02	0.036	0.013	0.063	0.027	0.013	0.041	1.3
SD, σ_τ	Inverse Gamma	[c]	[d]	0.196	0.178	0.216	0.036	0.033	0.040	5.5
SD, σ_k	Inverse Gamma	[c]	[d]	0.187	0.170	0.205	0.066	0.061	0.072	2.8
SD, σ_n	Inverse Gamma	[c]	[d]	0.157	0.143	0.172	0.035	0.032	0.038	4.5

Notes. [a] Uses the OLS estimate of the corresponding coefficient in the policy rule. [b] Uses 0.05×OLS estimate of the corresponding coefficient in the policy rule. [c] Uses the standard deviation of the residual from the estimated policy rule. [d] Uses 0.05×SD of the residual from the estimated policy rule.

research studies.[16] Also, the response of government expenditure to debt (θ_g^b) decreases by a factor of 2.8 from the pre- to the post-KW sample. In Section 7, we show that these changes seem to have produced only minor effects on the multiplier.[17]

Third, all standard deviations are smaller in the second period; of major importance is the moderation in the volatility of defense expenditure σ_m, which declines by a factor of 10. The standard deviations for the tax rates also experience a sensible moderation, shrinking by factors between 2.8 and 5.5 times;[18] these changes can be interpreted as a measure of fiscal discretion.[19] Figure 3 illustrates this finding. We look at the kernel densities for the posterior distributions of $\{\sigma_m, \sigma_{\text{war}}, \sigma_g, \sigma_{\text{gov}}\}$: solid lines correspond to the densities for the first sample, 1939–1954, while dashed lines are those of the second sample, 1955–2017.[20] In all cases, the dispersion in the posterior distribution falls during the second sample, which implies a reduction in the uncertainty associated with the expenditure shocks.

All together, the decline in the standard deviations of the fiscal rules, paired with an increase in their persistency (keeping the automatic stabilizers constant) point to a lower ability of U.S. authorities to conduct discretionary policy during the second sample.

6 Model Results

In this section, we first use the estimated parameters to compare our war news shocks—the innovations implied by our DSGE model—to the ones derived by Ramey and Zubairy (2018) under a narrative

[16] Leeper et al. (2010) offer posterior modal values ranging within $[0.12, 0.58]$ using quarterly data, while Afonso et al. (2010) estimate a value of 0.27 using annual data.

[17] Zubairy (2014) presents evidence that the multiplier falls as the responsiveness of government expenditure to debt increases.

[18] Fernández-Villaverde, Guerrón-Quintana, Kuester, and Rubio-Ramírez (2015) estimate fiscal rules for spending and tax rates with time varying volatility. They do not consider anticipated changes in government spending or the tax rates, though. They report evidence that these shocks to the time varying volatility of fiscal variables can produce adverse effects on economic activity.

[19] Afonso et al. (2010) provide evidence of a worldwide trade-off between persistence and discretion and, in turn, between persistence and responsiveness. Of course, their regressions should be seen as consistent correlations and not as causal relationships.

[20] To generate the kernels, we use the simulated parameters in the Metropolis-Hastings sampling algorithm, keeping the same number of values that are used in calculating the posterior means shown in Tables 2 and 4.

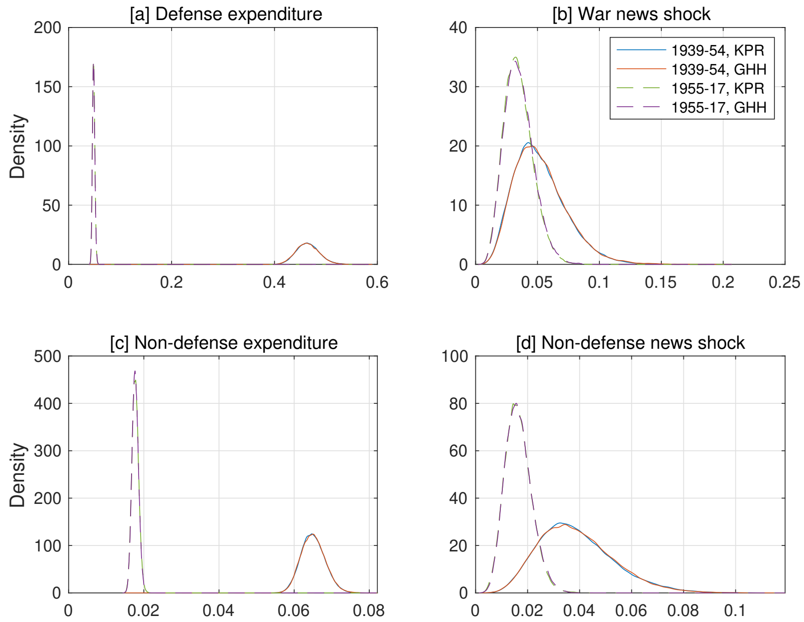

Figure 3: Kernel densities for posterior structural standard deviations.

approach. We then present a selection of impulse-response functions and comment on the model's ability to match U.S. business cycle statistics. Finally, we perform a variance decomposition exercise and close by presenting our estimates of expenditure multipliers.

6.1 Identifying the news of war

We use the smoothed shocks from the estimation procedure to identify the news of war implied by our model. Comparing these to the narrative defense news derived by Ramey (2011), we find that the correlation coefficients are 0.52 for the first sample and 0.36 for the second one when using KPR preferences, and 0.48 and 0.36 when using GHH preferences.

Figure 4 considers the pre-KW sample. The DSGE-based approach does a good job in capturing the dynamics of the defense news series found in Ramey and Zubairy (2018). While the DSGE-based shocks reproduce the ones from the narrative approach for WWII, our results suggest that the Korean War was largely unanticipated, which is consistent with the events that preceded the start of the war.[21]

Figure 5 presents the war news shocks for the post-KW sample. For the most part, both series move

[21] A bit of history: Secretary of State Dean Acheson, during his speech in the National Press Club in January 1950, did not consider the Korean Peninsula to be a part of the all-important "defense perimeter" of the U.S. The Korean War broke out shortly after, on June 25, 1950.

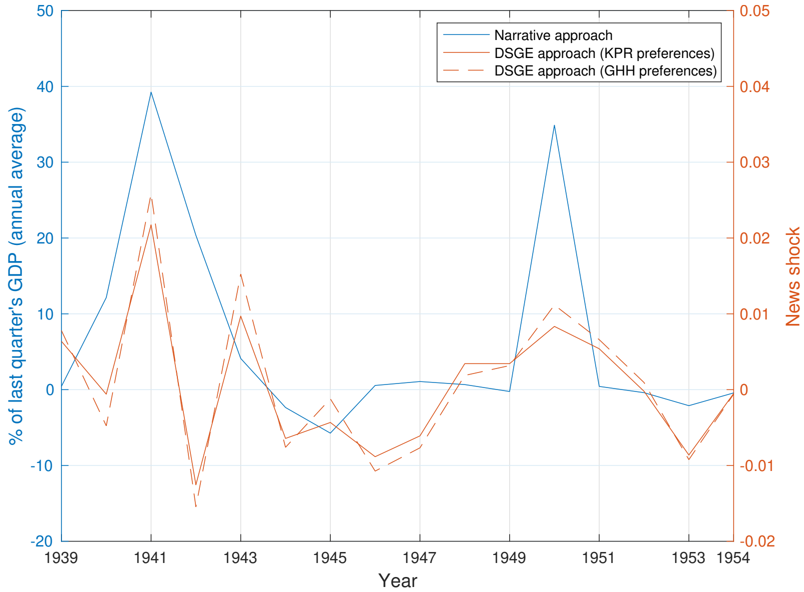

Figure 4: News of war, narrative vs. DSGE approaches (1939–1954).

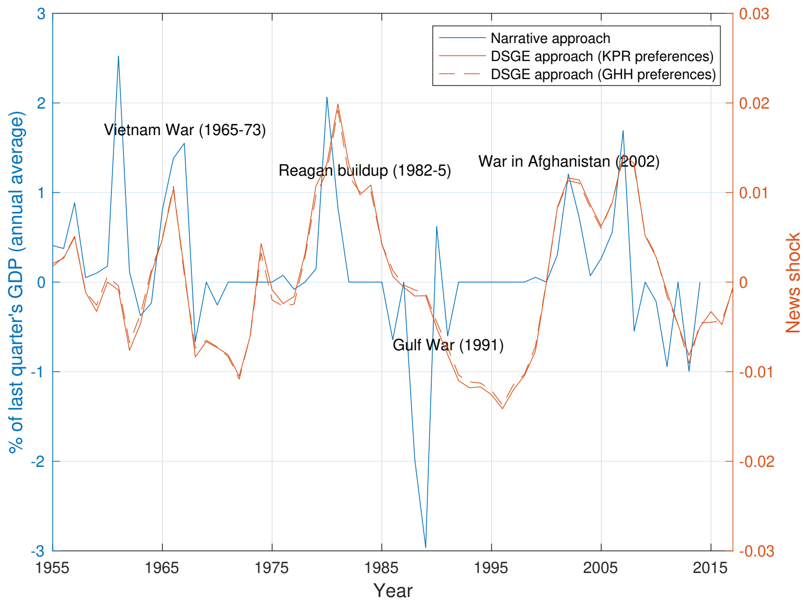

Figure 5: News of war, narrative vs. DSGE approaches (1955–2017).

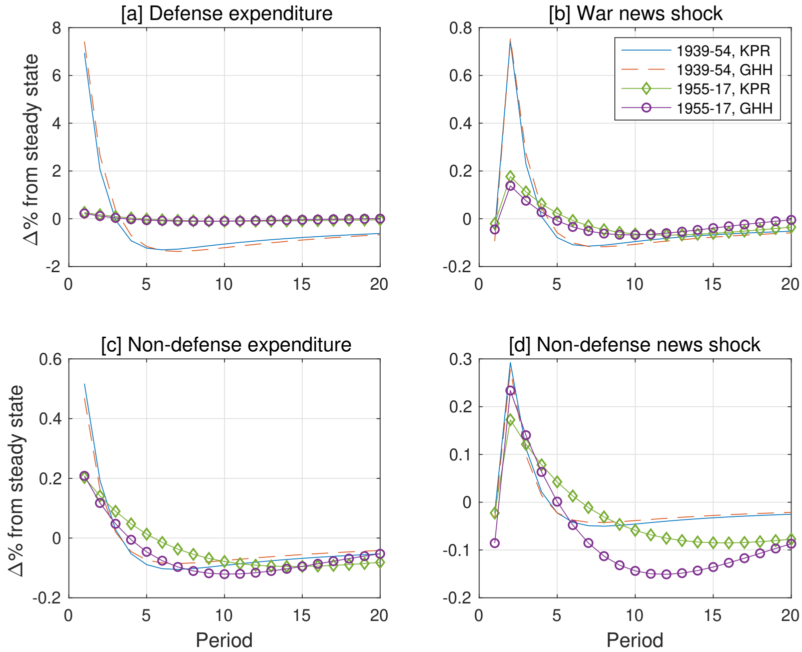

Figure 6: Model impulse-response functions (fiscal policy and news).

together, as is clear from the Vietnam War (1967), Reagan buildup (1982–5), and War in Afghanistan (2002) episodes. Ramey and Zubairy's narrative approach shows a contraction in defense expenses by the end of the 1980s: with the Cold War ending and nuclear arsenals shrinking, the 1990s were mostly peaceful, so changes in military spending can be taken as unanticipated during this decade.

6.2 Impulse-response functions

Figure 6 shows the impulse-response functions for output, following a one-standard deviation increase in defense expenditure, war news shocks, non-defense expenditure, and government expenditure news shocks. From panel [a], the non-anticipated defense expenditure shocks are fairly large in the pre-KW period (regardless of the preference specification we use) but flatten out after the Korean War occurs. The effect of the war news shock (panel b) is considerably smaller than the previous one and the magnitude of the change in output falls by a factor of 4 between the pre- and post-KW samples. Non-defense expenditure (panel c) exhibits an even smaller response to output, and the same can be said about the government expenditure news shock (panel d). Note that in all of these cases, the impulse-response functions moderate for 1955–2017, consistent with the results from Table 4.

More precisely, output increases on impact in response to unanticipated fiscal shocks (ε_{mt} and ε_{gt}, as shown in the left panels). Due to the negative wealth effect caused by the fiscal expansion, households reduce their consumption and work more hours. Both public debt and bond yields move up. These facts, combined with the arbitrage condition that comes from the model's first-order conditions, boost the real interest rate and reduce investment. After a few periods, the positive effects on output from the fiscal expansion are dampened by the fall in private expenditure. Note that output is more sensible to short-run fiscal shocks in the first sample, given the moderation in standard deviations (Table 4). A positive shock to the unanticipated defense spending shock ε_{mt} boosts military expenditure and readjusts the government budget internally, reducing non-defense disbursements by θ_m^g for any marginal dollar spent on defense needs above the steady state value m. Conversely, a shock to non-defense expenditure ε_{gt} has no impact on defense expenses. This asymmetrical adjustment results in a different response in output depending on the type of the fiscal shock.

Anticipated military expansions impact output differently. The war news shock $\varepsilon_{mt}^{\text{war}}$ has no effect on current non-defense expenditure, but does impact future outlays. Given the tax rules, families discount tax burden increases that may spread across a long horizon due to the possibility of war. As public debt is expected to rise, the real interest rate rises while investment falls. The net effect is an immediate contraction in output. In the second period, output reacts positively to the fiscal news, as the negative wealth effect causes households to demand less leisure or to work more hours, although private consumption and investment continue to fall. Eventually, the net response of output to the anticipated fiscal expansion becomes negative.[22]

6.3 Business cycle implications

We compare U.S. business cycle statistics with the ones derived from model simulations. Panel [a] in Table 5 presents the observed volatility of detrended series. The standard deviations fall drastically in most series, up to a factor of 4.2 (for defense spending). Also, there are remarkable changes in the correlation coefficients, as shown in panel [b]: for instance, the correlation of output with consumption more than doubles during the post-KW sample, while the output-investment correlation shifts from counter- to procyclical—consistent with a crowding-out effect during the first sample. While the correlation of non-

[22] Ramey (2011) finds a positive immediate response of output and services consumption to the defense news, which is contrary to our findings. To the extent that the participation of services on private consumption has been growing over time in the U.S., this suggests that her work underestimates the negative wealth effects associated with the shock.

Table 5: Business cycle facts of U.S. aggregate variables, 1939–2017.

Variable	[a] Standard deviation			[b] Correlation with y	
	1939–54	1955–17		1939–54	1955–17
Output, y	0.081	0.035		1.000	1.000
Consumption, c	0.024	0.023		0.369	0.941
Investment, x	0.187	0.082		-0.920	0.817
Non-defense expenditure, g	0.063	0.063		-0.638	0.697
Defense expenditure, m	0.509	0.122		0.903	0.035
Debt, b	0.194	0.109		0.404	-0.461
[c] Correlations of non-defense expenditure and investment with defense expenditure					
1939–54	-2	-1	0	1	2
$\text{corr}(g_t, m_{t+j})$	-0.437	-0.623	-0.502	0.279	0.687
$\text{corr}(x_t, m_{t+j})$	0.085	-0.557	-0.773	-0.144	0.268
1955–17					
$\text{corr}(g_t, m_{t+j})$	-0.084	-0.055	-0.031	-0.042	-0.061
$\text{corr}(x_t, m_{t+j})$	-0.423	-0.395	-0.324	-0.171	0.014

defense expenses with output changes from negative to positive between samples, debt does the opposite. Defense expenditure changes from highly procyclical to acyclical. Panel [c] presents the correlations of detrended non-defense outlays and investment with lags and leads of detrended defense spending. These confirm that both defense and non-defense expenditure are contemporaneously negatively correlated and that military disbursements are a leading indicator for non-defense spending.

Tables 6a (KPR preferences) and 6b (GHH) report the model-based moments. Comparing with Table 5, we conclude that our model—especially the GHH specification—can reproduce the moderation occurring after the Korean War. That said, while the model can accurately capture the fall in the volatility of defense spending, the adjustment factor in the moderation doesn't share the same magnitude as in the data. For instance, the model overestimates the moderation for investment, and implies that debt is five times more volatile than its sample analogue.

Regarding the correlations with output, the simulated values tend to meet the ones reported in Table 5. The GHH specification, in particular, captures the increase in the correlation between consumption and output, the change in cyclicality in investment, and the decrease in the correlation between defense expenditure and output. However, it cannot replicate the change in cyclicality in the correlation between non-defense disbursements and output, though it qualitatively reproduces the fall in the debt-to-output correlation.

Also, the model-based correlation between output and defense expenditure drops from 0.80 in the

Table 6a: Simulated moments for U.S. aggregate variables, 1939–2017, KPR preferences.

Variable	[a] Standard deviation			[b] Correlation with y	
	1939–54	1955–17		1939–54	1955–17
Output, y	0.087	0.026		1.000	1.000
Consumption, c	0.011	0.019		0.044	0.825
Investment, x	0.098	0.047		0.042	0.760
Non-defense expenditure, g	0.442	0.109		0.365	0.565
Defense expenditure, m	0.521	0.119		0.805	0.047
Debt, b	0.978	0.218		-0.378	-0.851
[c] Correlations of non-defense expenditure and investment with defense expenditure					
1939–1954	-2	-1	0	1	2
corr(g_t, m_{t+j})	-0.330	-0.294	-0.238	-0.112	-0.054
corr(x_t, m_{t+j})	-0.499	-0.525	-0.469	-0.199	-0.079
1955–2017	-2	-1	0	1	2
corr(g_t, m_{t+j})	-0.433	-0.416	-0.387	-0.342	-0.302
corr(x_t, m_{t+j})	-0.065	-0.069	-0.065	-0.049	-0.033

Table 6b: Simulated moments for U.S. aggregate variables, 1939–2017, GHH preferences.

Variable	[a] Standard deviation			[b] Correlation with y	
	1939–54	1955–17		1939–54	1955–17
Output, y	0.095	0.034		1.000	1.000
Consumption, c	0.026	0.026		0.613	0.933
Investment, x	0.114	0.057		-0.076	0.716
Non-defense expenditure, g	0.469	0.120		0.392	0.619
Defense expenditure, m	0.521	0.121		0.798	0.022
Debt, b	1.027	0.262		-0.388	-0.895
[c] Correlations of non-defense expenditure and investment with defense expenditure					
1939–1954	-2	-1	0	1	2
corr(g_t, m_{t+j})	-0.333	-0.289	-0.230	-0.109	-0.053
corr(x_t, m_{t+j})	-0.551	-0.583	-0.517	-0.216	-0.081
1955–2017	-2	-1	0	1	2
corr(g_t, m_{t+j})	-0.406	-0.390	-0.364	-0.324	-0.289
corr(x_t, m_{t+j})	-0.065	-0.073	-0.072	-0.057	-0.041

first sample to 0.02 in the second one. In Section 3, we saw that the slope of (3.1) can be interpreted as an estimate of the fiscal multiplier: $m_Y = \text{cov}(\Delta Y_t/Y_t, \Delta M_t/Y_t)/\text{var}(\Delta M_t/Y_t)$. Thus, it follows that the model can account for the downturn in the slope coefficient m_Y for the second sample, as reported in Figure 2.

Finally, the simulated correlations of non-defense outlays and investment with leads and lags of defense expenditure replicate the leading indicator feature from Table 5: defense outlays are a leading indicator of investment in the first sample, yet the model implies an orthogonal relation between investment and defense expenditure at all leads and lags in the second sample.

6.4 Variance decomposition analysis

Tables 7a (KPR preferences) and 7b (GHH) show the variance decomposition for the main aggregate variables. (Due to space limitations, we omit the contributions to the shocks to transfers and taxes. We also summarize the sum of the contributions of the non-defense news and war news shocks by η, where $\eta \equiv \varepsilon_{mt}^{\text{war}} + \varepsilon_{gt}^{\text{gov}}$.) Since we work with annual series, we present an analysis of conditional variances on impact (at $j = 0$) and a horizon of up to 4 years.

We start by describing the results that hold for both sets of preferences. First, in the pre-KW sample, defense fiscal shocks have an important effect on output, accounting for 99% of output variance on impact and 95% five years ahead. Fiscal shocks from discretionary non-defense policies (ε_{gt}) play a minor role. The bulk of the variability in consumption, investment, and hours worked is mostly accounted for by unanticipated defense expenditure shocks (ε_{mt}). Moreover, unanticipated fiscal spending shocks account for nearly all of the short run variability of debt. Surprisingly, anticipated shocks (both defense and non-defense) have a negligible impact over the variability of model variables.

For the post-KW sample, variability in output, consumption, investment, government debt, and hours worked is mostly accounted for by TFP shocks. Unanticipated expenditure shocks are unimportant, each accounting for 5% or less of the variability in output on impact. Since the variance of defense expenditure falls considerably in the second sample, the unanticipated defense shock reduces its importance for non-defense spending. The war news shock matters for defense expenditure (at a 5-year horizon), accounting for 32% of its variance; non-defense news matter for non-defense spending, accounting for 47% of the variability. The role of anticipation on the remaining variables is virtually zero.

Some differences between Tables 7a and 7b are worth pointing out. First, in the pre-KW sample,

Table 7a: Decomposition of simulated conditional variances, KPR preferences.

Variable	Year	1939–54				1955–17			
		ε_z	ε_m	ε_g	η	ε_z	ε_m	ε_g	η
Output	0	0.01	0.99	0.00	0.00	0.89	0.05	0.05	0.00
	4	0.04	0.94	0.00	0.01	0.94	0.02	0.02	0.03
Consumption	0	0.49	0.50	0.00	0.00	0.99	0.00	0.00	0.00
	4	0.51	0.41	0.00	0.02	0.96	0.00	0.01	0.02
Investment	0	0.38	0.61	0.00	0.00	0.98	0.01	0.01	0.01
	4	0.30	0.68	0.00	0.01	0.92	0.01	0.02	0.04
Non-defense expenditure	0	0.00	0.46	0.54	0.00	0.00	0.04	0.96	0.00
	4	0.08	0.78	0.08	0.04	0.07	0.15	0.27	0.51
Defense expenditure	0	0.00	1.00	0.00	0.00	0.00	1.00	0.00	0.00
	4	0.00	0.99	0.00	0.01	0.00	0.68	0.00	0.32
Debt	0	0.02	0.86	0.00	0.00	0.51	0.09	0.09	0.00
	4	0.17	0.77	0.00	0.00	0.85	0.03	0.04	0.03
Hours worked	0	0.22	0.78	0.00	0.00	0.96	0.01	0.01	0.00
	4	0.21	0.77	0.00	0.01	0.91	0.02	0.02	0.03

Table 7b: Decomposition of simulated conditional variances, GHH preferences.

Variable	Year	1939–54				1955–17			
		ε_z	ε_m	ε_g	η	ε_z	ε_m	ε_g	η
Output	0	0.00	0.99	0.00	0.00	0.88	0.05	0.04	0.01
	4	0.03	0.95	0.00	0.01	0.95	0.01	0.01	0.02
Consumption	0	0.04	0.92	0.00	0.01	0.94	0.00	0.00	0.01
	4	0.03	0.92	0.00	0.00	0.94	0.00	0.00	0.01
Investment	0	0.32	0.68	0.00	0.00	0.97	0.01	0.01	0.01
	4	0.25	0.73	0.00	0.01	0.91	0.01	0.02	0.05
Non-defense expenditure	0	0.00	0.46	0.54	0.00	0.00	0.04	0.96	0.00
	4	0.08	0.79	0.07	0.03	0.07	0.14	0.26	0.51
Defense expenditure	0	0.00	1.00	0.00	0.00	0.00	1.00	0.00	0.00
	4	0.00	0.99	0.00	0.01	0.00	0.68	0.00	0.32
Debt	0	0.02	0.86	0.00	0.00	0.59	0.09	0.09	0.00
	4	0.16	0.79	0.00	0.00	0.87	0.03	0.04	0.03
Hours worked	0	0.14	0.85	0.00	0.00	0.14	0.07	0.06	0.07
	4	0.14	0.84	0.00	0.01	0.61	0.03	0.02	0.06

Note. η denotes the sum of the contributions of the non-defense news and war news shocks.

consumption is equally explained by TFP and unanticipated defense shocks in the KPR case, but fully accounted for by unanticipated defense shocks under GHH preferences. Because the latter set of preferences doesn't incorporate wealth effects, TFP shocks play a trivial role in accounting for the behavior of consumption. Similarly, in the post-KW sample, the variability of hours worked under KPR preferences is almost entirely accounted for by TFP shocks (96%), yet under GHH preferences we find that labor income tax shocks are the primary driver of changes in this variable—at least on impact (though this is not shown in the table, the corresponding value is 64%). The explanation is similar: under GHH preferences, a shock to labor income taxes induces a pure substitution effect, affecting the choice of hours worked.

Our results are contrary to those of Ben Zeev and Pappa (2017); using a VAR and quarterly data from 1948 to 2007, they find that news shocks account for 9% and 13% of variation in output and hours at a one-year horizon. Under the best-case scenario (post-KW sample at a 5-year horizon), the equivalent values in our analysis are 3% for output (assuming KPR preferences) and 7% for hours worked (under GHH preferences). Note that these values include the effect of both news shocks.[23]

6.5 Measuring expenditure multipliers

Let Δy_{t+s} and Δe_{t+s}, $s \in \{0, 1, 2, \ldots\}$, denote the impulse-response functions of output and public expenditure with respect to fiscal shock ε_{et}, where $e = \{g, m\}$. Following Mountford and Uhlig (2009), the present-value multiplier generated by a change in e_t over a j-period horizon is

$$\mathrm{PV}_e^y(j) = \frac{\sum_{s=0}^{j} \Delta y_{t+s} \times \prod_{s=0}^{j} (r_{t+s}^b)^{-1}}{\sum_{s=0}^{j} \Delta e_{t+s} \times \prod_{s=0}^{j} (r_{t+s}^b)^{-1}} \cdot \frac{y}{e}, \tag{6.1}$$

where r_{t+s}^b is the impulse-response function for the government bond yield. The multipliers for consumption, investment, non-defense, and defense expenditure are defined accordingly.

Tables 8a (KPR preferences) and 8b (GHH) show the present value multipliers implied by our model, up to an horizon of 4 years. (For time horizon $j = 0$, the multiplier is measured on impact.) Because these multipliers are calculated using the impulse-response functions derived from our structural DSGE model, the estimate of the multiplier is immune to identification bias. When considering the multiplier

[23] Ramey (2011) uses a sample similar to the one in Ben Zeev and Pappa (2017) and finds that these shares are 2% (ouptut) and 4% (hours).

Table 8a: Present value multipliers, KPR preferences.

Sample	Output PV_g^y	PV_m^y	Consumption PV_g^c	PV_m^c	Investment PV_g^x	PV_m^x
1939–54						
$j=0$	0.920	0.869	-0.010	-0.009	-0.070	-0.067
1	0.870	0.790	-0.015	-0.013	-0.115	-0.104
2	0.816	0.705	-0.019	-0.017	-0.165	-0.141
3	0.758	0.617	-0.023	-0.019	-0.219	-0.175
4	0.697	0.532	-0.026	-0.020	-0.277	-0.206
1955–17						
$j=0$	0.755	0.648	-0.124	-0.098	-0.121	-0.092
1	0.668	0.522	-0.155	-0.112	-0.177	-0.122
2	0.592	0.416	-0.177	-0.115	-0.231	-0.145
3	0.522	0.326	-0.193	-0.112	-0.285	-0.162
4	0.454	0.247	-0.207	-0.106	-0.340	-0.173

Table 8b: Present value multipliers, GHH preferences.

Sample	Output PV_g^y	PV_m^y	Consumption PV_g^c	PV_m^c	Investment PV_g^x	PV_m^x
1939–54						
$j=0$	0.972	0.917	0.062	0.058	-0.091	-0.087
1	0.944	0.855	0.094	0.083	-0.150	-0.136
2	0.908	0.779	0.124	0.103	-0.216	-0.184
3	0.862	0.694	0.151	0.117	-0.289	-0.228
4	0.807	0.605	0.174	0.124	-0.367	-0.267
1955–17						
$j=0$	0.767	0.672	-0.060	-0.040	-0.173	-0.129
1	0.668	0.538	-0.083	-0.052	-0.249	-0.170
2	0.572	0.421	-0.106	-0.061	-0.322	-0.199
3	0.478	0.317	-0.131	-0.070	-0.391	-0.217
4	0.382	0.227	-0.159	-0.077	-0.459	-0.227

implied by defense expenditure, $e = m$, the fiscal adjustment taking place within government budget—parameter θ_g^m in equation (4.12)—is internalized in the model: a positive fiscal shock ε_{mt} boosts defense spending over its steady value m, prompting the government to adjust non-defense outlays by $-\theta_g^m$.

We draw four conclusions from Table 8. First, the output multipliers are always below unity and the investment multipliers are negative, an indication of crowding out.[24] Moreover, both the defense and non-defense multipliers decline (in absolute value) as we compare pre- and post-KW values. Second, non-defense expenditure multipliers are slightly higher than defense ones, a result from the asymmetric budgetary adjustment motivated by θ_g^m. Third, the multipliers that arise from a GHH preference specification are slightly larger than those coming from KPR preferences, explained mostly by a positive consumption response to the fiscal expenditure shocks in the GHH case (investment falls more in the GHH case relative to the KPR scenario, but the total effect results in larger multipliers in the former case). Finally, the value of the multiplier decreases as the time horizon increases.

Distribution of the multiplier Figures 7 (KPR) and 8 (GHH) show the distribution of the multipliers for both samples.[25] As is clear from the left panels in both figures, the distribution of the post-KW output multiplier rarely intersects with that of the pre-KW one. While the consumption (center panel) and investment multipliers (right panel) exhibit larger overlaps, our central message remains: post-KW multipliers are smaller.

7 Why Did the Magnitude Decline?

Finally, we present a counterfactual experiment for the PV output multipliers, in order to understand what can account for the fall in their magnitude. We consider the posterior means given in Tables 3 and 4; these parameters are classified into three groups: (1) private-agent parameters, which include household and firm parameters; (2) environment parameters, namely, the steady state ratios of debt and transfers and the average GDP composition; and (3) fiscal parameters.

[24] The inability of the neoclassical framework to produce large multipliers has been highlighted by Dyrda and Ríos-Rull (2012) and Ríos-Rull and Huo (2013), who suggest introducing other frictions to motivate higher multipliers.

[25] We follow the same procedure as in Figure 3: use the simulated parameters in the Metropolis-Hastings sampling algorithm, keeping the same number of values that are used in calculating the posterior mean values, and then calculate the associated multiplier for each of these cases.

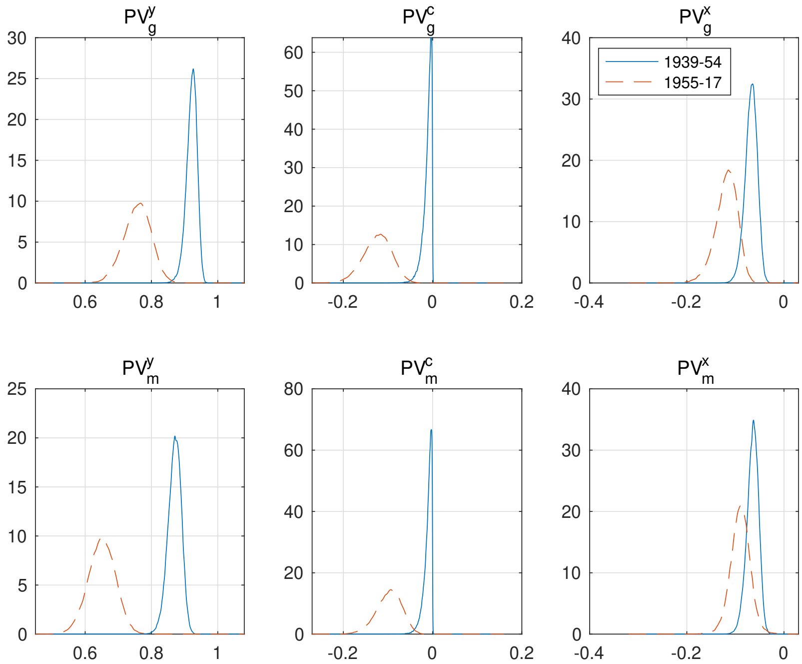

Figure 7: Distribution of the multiplier, KPR preferences.
(Solid line: 1939–1954. Dashed line: 1955–2017.)

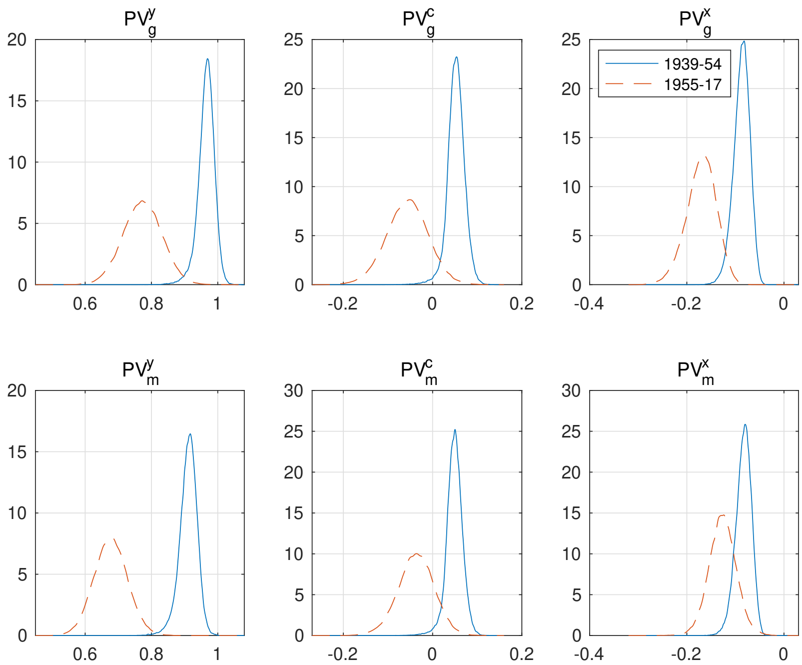

Figure 8: Distribution of the multiplier, GHH preferences.
(Solid line: 1939–1954. Dashed line: 1955–2017.)

To put things in context, the first two rows in Table 9a reproduce the estimates of the output multipliers on impact; as discussed earlier, the post-KW multipliers fall by factors that range between 1.2 and 1.4, which implies that government's ability to affect output using spending policies is lower after the Korean War. The last row presents the estimated change in these multipliers.

Table 9b reports the change in the impact post-KW output multipliers when changing one parameter at a time: we take the parameter set of the second sample as a benchmark and impose the corresponding values of the first sample, one by one. Once the "new post-KW" (counterfactual) multipliers have been computed, we present the predicted change in the multiplier. To avoid unnecessary clutter, we only list parameters that generate changes over 0.01 (1 cent).

Clearly, the bulk of the change in the PV multipliers is accounted by a change in household parameters and in fiscal rules parameters, groups (1a) and (3a), respectively. More concretely, the fall in the consumption habit persistence parameter μ (from 0.85 to 0.53) can account for 7.3 and 8.8 cents in the KPR defense and non-defense PV multipliers, respectively. Analogously for GHH preferences, consumption habit persistence can explain a change of 6.9 and 8.0 cents in the defense and non-defense PV multipliers. In both cases, these values represent between a quarter and a half of the estimated change in the PV multipliers. As we highlighted in Table 2, the estimated value of μ decreased in the post-KW sample, which favors a (more) negative wealth effect response in private consumption and, consequently, a fall in the multiplier.

To a lesser extent, the intertemporal elasticity of substitution γ (which in the second sample decreased from 1.98 to 1.57 in the KPR case, or from 1.61 to 1.57 in the GHH case) explains a minor fraction in this change: lower risk aversion implies a higher intertemporal elasticity of substitution, which exacerbates the wealth effect. The remaining private agent parameters, as well as the parameters related to the economy environment, do not influence the estimated change of the multipliers.

The PV multipliers are also sensitive to the autocorrelation coefficients of public expenditure, ρ_g and ρ_m. As shown in Table 4, the persistency in the fiscal rules increases in the second sample: from 0.42 to 0.82 for ρ_g and from 0.45 to 0.89 for ρ_m.[26] For non-defense expenditure multipliers, the increased persistency entails a fall in the multiplier of 6.4 and 11.7 cents for the KPR and GHH cases, respectively. These account for 40 and 57% of the total change. The rise in fiscal persistence after the Korean War

[26] As seen in Section 5.2, these estimates are consistent with those reported by other authors using similar data and time span (Afonso et al. 2010).

Table 9a: Change in present value multipliers.

Sample	KPR preferences PV_y^g	KPR preferences PV_y^m	GHH preferences PV_y^g	GHH preferences PV_y^m
1939–1954	0.920	0.869	0.972	0.917
1955–2017	0.755	0.648	0.767	0.672
Estimated change	0.165	0.221	0.205	0.245

Table 9b: Change in PV multiplier when switching. . .

Parameter(s)	KPR preferences ΔPV_y^g	KPR preferences ΔPV_y^m	GHH preferences ΔPV_y^g	GHH preferences ΔPV_y^m
(1a) All household parameters (μ, χ, γ)	0.085	0.069	0.068	0.058
Only habit formation (μ)	0.088	0.073	0.080	0.069
Only the inverse of the Frisch elasticity (χ)	-0.013	-0.011	-0.024	-0.022
Only the inverse elasticity of substitution (γ)	0.021	0.016	0.003	0.002
(1b) All firm and technology parameters $(\delta_0, \delta_2, \kappa, \rho_z, \sigma_z)$	0.002	0.007	-0.006	-0.001
Only steady-state depreciation term (δ_0)	0.002	0.007	0.014	0.015
Only quadratic depreciation term (δ_2)	0.000	0.000	-0.020	-0.017
(2) All environment ratios (b, τ, c, x, g, m)	0.000	0.068	-0.006	-0.001
GDP shares (c, x, g, m)	-0.007	0.060	0.000	0.000
(3a) All fiscal persistency parameters	0.069	0.047	0.122	0.084
Only military expenditure persistence (ρ_m)	0.000	0.058	0.000	0.105
Only government expenditure persistence (ρ_g)	0.064	-0.013	0.117	-0.028
(3b) All fiscal rule parameters (of the form θ_j^h)	0.017	-0.016	0.027	-0.002
Only non-defense response to military spending (θ_g^m)	0.000	-0.025	0.000	-0.020
Only non-defense response to debt (θ_g^b)	0.015	0.014	0.028	0.024
(3c) All structural standard deviations (of the form σ_j)	0.000	0.000	0.000	0.000
(3d) All average tax rates	-0.013	0.000	0.000	0.000
Total explained change (includes unreported parameters)	0.161	0.175	0.205	0.138
Unexplained change	0.004	0.046	0.000	0.107

reduced discretion and increased the wealth effect on consumption, moderating the spending multipliers.

We should add that these findings are consistent with other research studies, such as Fatás and Mihov (2006) or Afonso et al. (2010). The latter authors find that tighter fiscal constraints (higher ρ_m and ρ_g) lead to less discretion (lower σ_g and σ_m) and, consequently, less responsiveness of fiscal policies to output fluctuations (a decline in θ_g^y). More precisely, lower discretion (alternatively, higher persistence) can generate an ambiguous overall effect on macroeconomic stability. On one hand, lower discretion limits the ability of the fiscal authority to respond when negative shocks hit, as automatic stabilizers now play a lesser role. On the other hand, a smaller variance of the fiscal rules entails a lower output volatility. The model—especially the GHH specification—can reasonably reproduce the moderation occurring after the Korean War (Table 5 vs. Table 6). Following Fatás and Mihov (2006), tighter fiscal constraints after the Korean War led to increasing output stability, though the smaller multipliers limited the government's ability to affect output through fiscal policy.

Discussion In summary, the fall in the magnitude of the PV multipliers in the post-KW sample can be accounted by lower consumption habit persistence, paired with larger autocorrelation coefficients of public expenditure; as we have argued earlier, both of these changes lead to a stronger negative wealth effects on consumption. Overall, the differences due to preference specifications (KPR vs. GHH) are negligible.[27]

For completeness, Figures 9 to 12 present a sensitivity analysis where we include a wider range of values for $\{\mu, \chi, \gamma; \rho_g, \rho_m; \theta_g^m, \theta_g^y, \theta_g^b\}$—for which the PV multipliers are sensitive. As in Table 10, we take the parameters estimated for the second sample for granted and compute (ceteris paribus) the output PV multipliers changing the value of these parameters, one at a time. In all figures, shaded areas indicate 95% confidence bands around the post-KW posterior mean.

Figures 9 and 10 present the multiplier sensitivity with respect to household parameters. Consistent with Table 9, the change in these parameters can account for almost a half of the decline in the multiplier. Figures 11 and 12 present analogous results with respect to fiscal policy parameters $\{\rho_g, \rho_m; \theta_g^m, \theta_g^b, \theta_g^y\}$. For persistency parameters ρ_g and ρ_m, the posterior means of the pre-KW sample lie well outside the

[27] More importantly, the moderation in the structural standard deviations of the expenditure shocks (σ_g and σ_m; see Figure 3) had no effect on the PV multipliers: as highlighted in Section 6.5 and equation (6.1), the PV multiplier is neutral to changes in the scale of the structural variances.

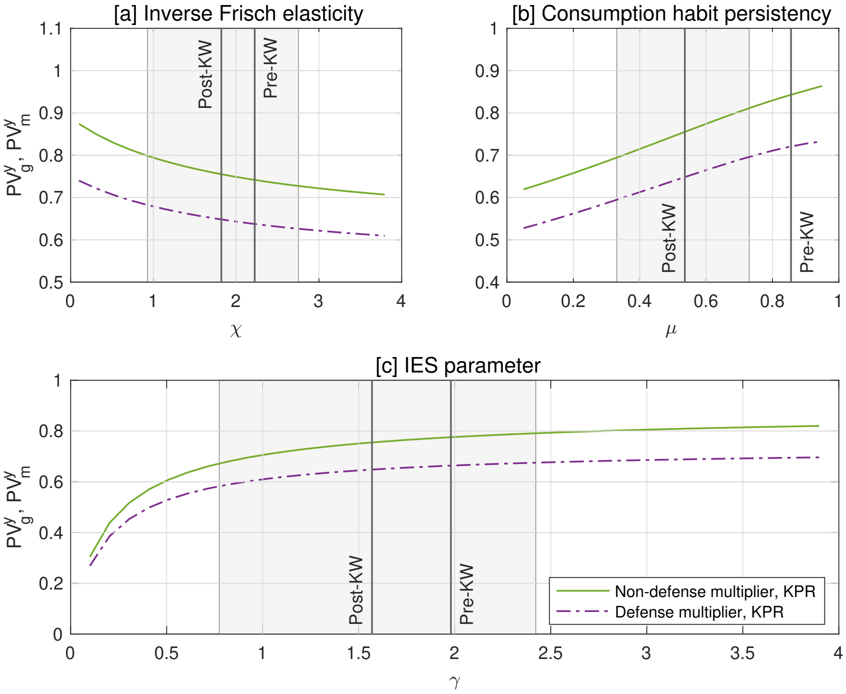

Figure 9: Sensitivity analysis of the impact PV multiplier, KPR preferences.

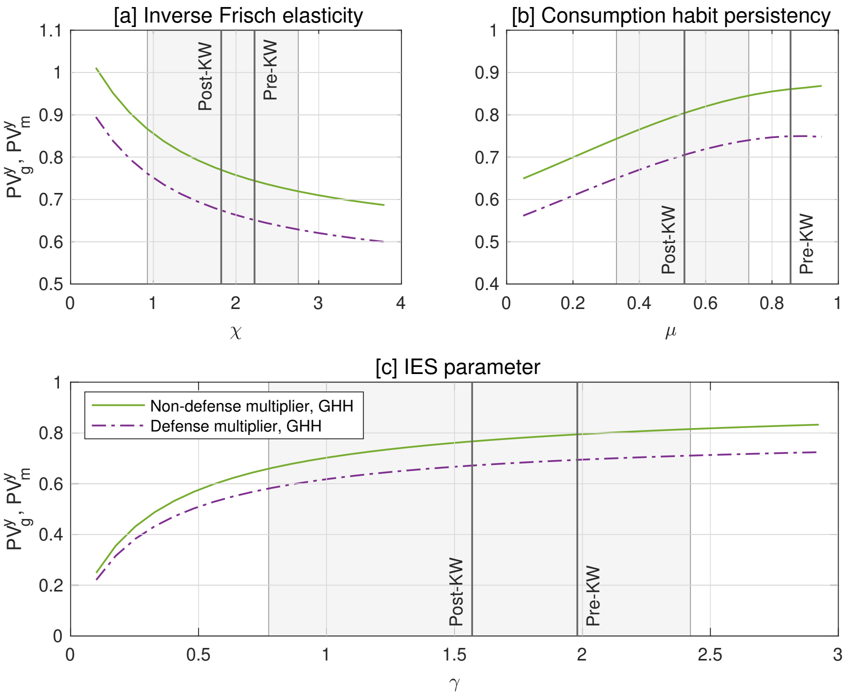

Figure 10: Sensitivity analysis of the impact PV multiplier, GHH preferences.

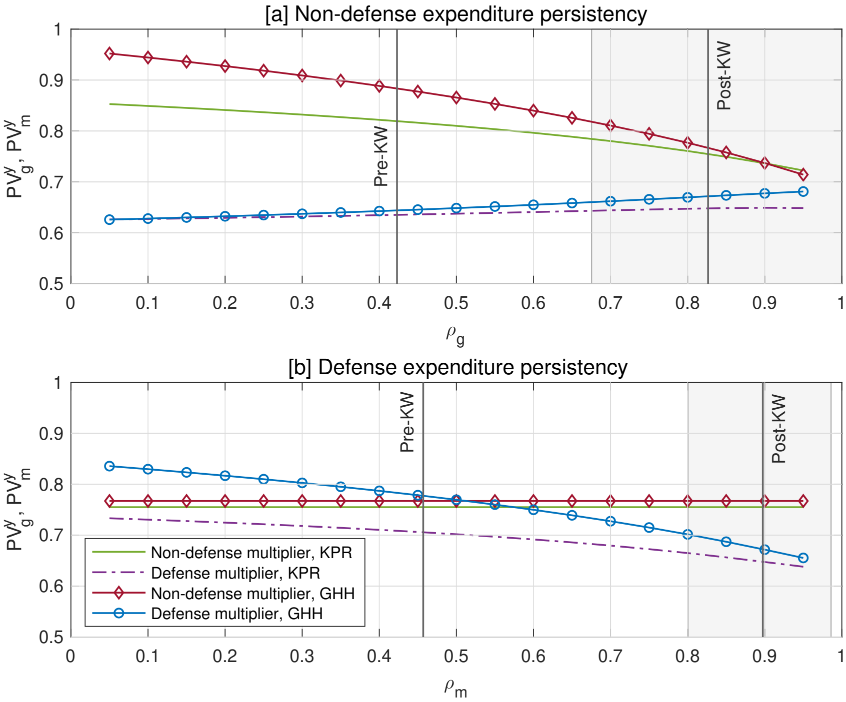

Figure 11: Sensitivity analysis of the impact PV multiplier, fiscal rule autocorrelations.

shadowed area 95% confidence band. Given our earlier discussion, the increase in the value of these parameters (which happens at the same time that structural variances decline) helps explain the decrease of fiscal discretion and the decline of the multipliers.

Concerning the remaining fiscal parameters $\{\theta^m_g, \theta^b_g, \theta^y_g\}$, panel [a] in Figure 12 shows a high sensitivity of the non-defense multiplier with respect to the internal budgetary adjustment: the lower θ^m_g, the higher the multiplier response. The slope $\Delta PV^y_m/\Delta\theta^m_g$ equals -1.14 under KPR preferences and -0.83 under a GHH specification, while the equivalent for non-defense spending, $\Delta PV^y_m/\Delta\theta^m_g$, remains flat. In other words, the lower the U.S. spending elasticity with respect to military outlays (θ^m_g), the higher effectiveness of fiscal policy is. To the best of our knowledge, this channel of internal adjustment has not been explored before.

Finally, panels [b] and [c] in Figure 12 show that the non-defense expenditure multipliers are flat with respect to the debt and output response parameters. This contrasts with the results of Leeper et al. (2010), Ilzetzki et al. (2013), and Zubairy (2014), who provide evidence that the multipliers are remarkably sensible to θ^b_g and θ^y_g.

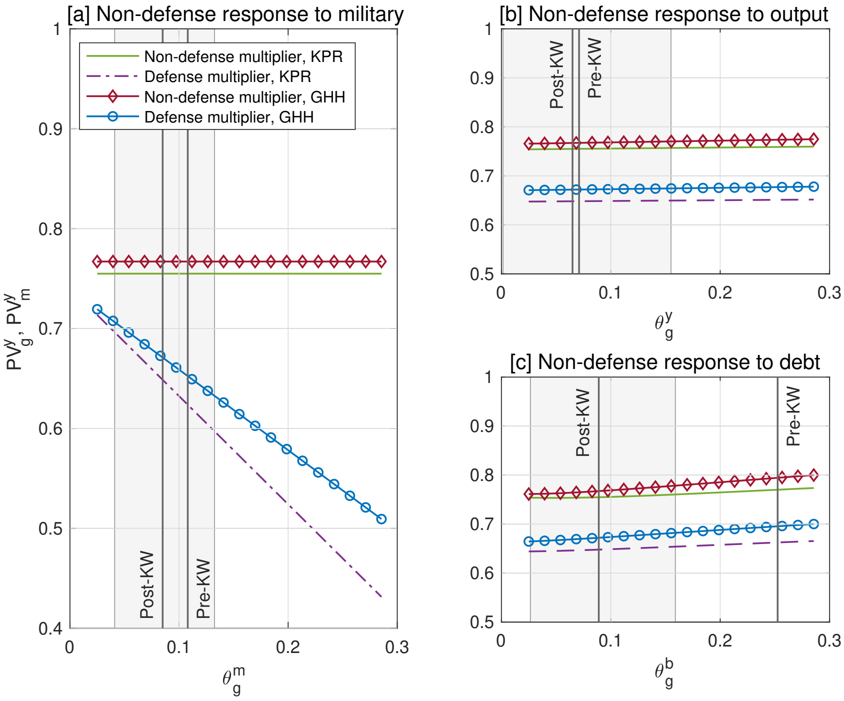

Figure 12: Sensitivity analysis of the impact PV multiplier, non-defense responsiveness.

8 Concluding Remarks

In this paper, we have documented a permanent fall in the magnitude of the U.S. multiplier that occurred after the Korean War. To analyze this decline, we have estimated a DSGE model using Bayesian techniques and a long-run data series. We conclude that the effectiveness of spending policies—as measured by the expenditure multiplier—has decreased in the post-KW sample due to changes in preferences and policy choices—from 0.97 to 0.77 in non-defense and 0.92 to 0.67 in defense spending (assuming GHH preferences).

As explained earlier, the decline in the value of the multiplier can be linked to changes in two model parameters: a fall in the consumption habit persistence (μ), paired with a rise in the persistence of the public expenditure processes (ρ_m and ρ_g). Other fiscal parameters—which earlier research papers have shown to be relevant, such as those linking expenditure with the state of debt (θ_b^g) or automatic stabilizers (θ_y^g)—have not affected the magnitude of the fiscal multiplier in the U.S. That said, the process by which spending is distributed among military and non-military outlays can be important. The role of the budgetary adjustment parameter (θ_m^g) might be relevant, though the observed decline in the multiplier should not be associated with it.

Complementarily, we have also documented a decline in the variability of news about future fiscal expenditure; this is, fiscal policies are now better anticipated than during 1939–1955, a period that involved two important warfare episodes. The pattern of fiscal news, on the other hand, has been adequately identified by the model, as compared with the narrative series proposed by Ramey (2011).

The debate about the effectiveness of fiscal policy has increased over the last decade, especially in the aftermath of the Great Recession and, almost a decade later, the Covid-19 crisis. Although the origins of these episodes are different, the usefulness of fiscal tools to cope with these crises depends on their ability to affect output. Under the lens of our model, these changes in fiscal scenarios have reduced the degree of discretion and, thereby, increased the role of fiscal rules. Put differently, our conclusion points to a limited role of discretionary increases of government spending to boost output at all time horizons.

A Appendix

A.1 Data and sources

The estimation procedure considers nine observable variables. We include consumption; investment; defense and non-defense government expenditure; transfers; tax rates on capital income, labor income, and consumption; and debt outstanding.

The first four series come from the Bureau of Economic Analysis (BEA). Since we want to incorporate observations that include WWII, our analysis is limited to annual series; quarterly series are provided by the BEA from 1947 to present only.

- Consumption (C) is defined as the sum of non-durable goods consumption and services consumption, $CNonDur + CServ$ (Table 1.1.5, lines 5 and 6).
- Investment (X) is defined as the sum of durable goods consumption and fixed investment, $CDur + FI$ (Table 1.1.5, lines 4 and 8).
- Total government expenditure consists of (federal-level) defense government expenditure and (federal, state, and local) non-defense government expenditure. In turn, defense government expenditure (M) is retrieved from Table 1.1.5, line 24, while non-defense expenditure (G) consists of federal spending plus state and local public expenditure (Table 1.1.5, lines 25 and 26).
- Transfers (τ) are obtained from the series `A577RC1A027NBEA`, personal current transfer receipts, from the St. Louis Fed's FRED database.
- Regarding consumption, labor income, and capital income tax rates, we follow the methodology of Mendoza, Razin, and Tesar (1994), who derive aggregate estimates for effective tax rates using national accounts. (Both Leeper et al. 2010 and Fernández-Villaverde et al. 2015 calculate tax rates with a similar strategy.) The consumption tax rate (τ_c) is computed as the sum of Excise taxes, $ExcT$, Customs duties, $CustDut$, and Sales taxes, $SalesT$ (Table 3.2, lines 5, 6, and 7), divided by the sum of Personal consumption expenditures C (Table 1.1.5, line 2) and government consumption expenditures $G + M$ (Table 3.9.5, line 2):

$$\tau_c = \frac{ExcT + CustDut + SalesT}{C + (G + M)}.$$

The income tax rate τ_Y—which is not used directly as an observable variable—is calculated as the ratio of Income tax revenues $IncT$ (Table 3.4, lines 3 and 9) over the sum of Proprietors' income with inventory valuation and capital consumption adjustments (PI, Table 2.1, line 9), Rental income of persons with capital consumption adjustment ($RICCA$, Table 2.1, line 12), Personal income receipts on assets ($PIRA$, Table 2.1, line 13), Wages and salaries (W), and Supplements to wages and salaries (SW) (Table 2.1, lines 3 and 6):

$$\tau_Y = \frac{IncT}{PI + RICCA + PIRA + (W + SW)}.$$

Using the calculated tax rate on income τ_Y, the tax rate on labor income (τ_n) is

$$\tau_n = \left(\tau_Y + \frac{GSI}{W + SW}\right)\frac{W + SW}{CE},$$

where GSI denotes Contributions for government social insurance (Table 3.1, line 7), and CE denotes Compensation of employees (Table 2.1, line 2). Finally, the tax rate on capital income τ_k) is computed as

$$\tau_k = \frac{\tau_Y(PI + RICCA + PIRA) + PT + TCI}{PI + PIRA + IRA},$$

where PT denotes Property taxes (Table 3.4, line 11), TCI denotes taxes on corporate income (Table 3.2, line 7; Table 3.3, line 10), and IRA denotes Income receipts on assets (Table 4.1, line 9).

- The dollar amount of debt b is obtained from the *Historical Debt Outstanding* statistics found in the Treasury Direct website.

All series (except for tax rates) are transformed into real series by dividing by the GDP implicit deflator (Table 1.1.4, line 1). Once we do this, all series described above are logged and regressed against a constant and a time trend. The residuals are the observable variables we use in the econometric analysis.

Calculated tax rates Figure 13 presents our estimates for the [a] labor, [b] capital, and [c] consumption tax rates, as well as [d] the debt-to-output ratio.

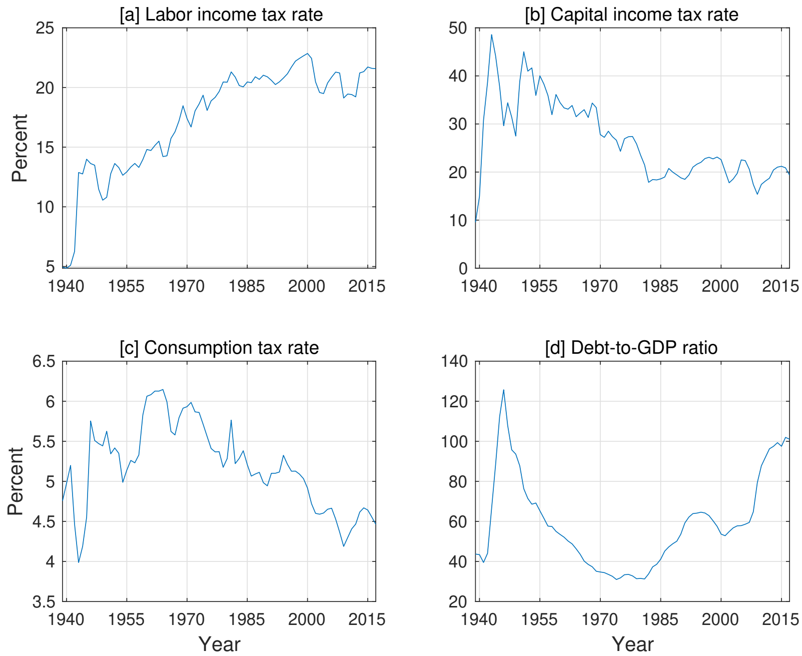

Figure 13: Selected fiscal variables, 1939–2017.

A.2 Equilibrium conditions

Given the model economy described in the paper, let λ_t denote the Lagrange multiplier on the budget constraint that results from merging equations (3.2) and (3.3) and ξ_t be the multiplier associated with the law of motion for capital (3.4). Then, the first-order conditions that characterize the economy's equilibrium are, for KPR preferences,

$$\lambda_t(1+\tau_{ct}) = q_t^{-1} - \beta\mu E_t q_{t+1}^{-1} \tag{A.1}$$

$$\begin{aligned}\lambda_t &= \xi_t - \xi_t S'\left(\frac{x_t}{x_{t-1}}\right)\frac{x_t}{x_{t-1}} - \xi_t S\left(\frac{x_t}{x_{t-1}}\right) \\ &\quad + \beta E_t \xi_{t+1} S'\left(\frac{x_{t+1}}{x_t}\right)\left(\frac{x_{t+1}}{x_t}\right)^2\end{aligned} \tag{A.2}$$

$$\lambda_t(1-\tau_{nt})w_t = \phi n_t^{\chi} \tag{A.3}$$

$$\lambda_t(1-\tau_{kt})r_t = -\lambda_t \delta'(u_t)\tau_{kt} + \xi_t \delta'(u_t) \tag{A.4}$$

$$\begin{aligned}\xi_t &= \beta E_t \lambda_{t+1}(1-\tau_{k,t+1}) r_{t+1} u_{t+1} \\ &\quad + \beta E_t \lambda_{t+1} \delta(u_{t+1}) \tau_{k,t+1} + \beta E_t \xi_{t+1} \\ &\quad - \beta E_t \xi_{t+1} \delta(u_{t+1})\end{aligned} \tag{A.5}$$

$$\lambda_t = \beta E_t \lambda_{t+1} r_t^b \tag{A.6}$$

$$k_{t+1} = k_t - \delta(u_t)k_t + x_t - x_t S\left(\frac{x_t}{x_{t-1}}\right) \tag{A.7}$$

$$y_t = z_t(u_t k_t)^{\alpha} n_t^{1-\alpha} \tag{A.8}$$

$$\alpha y_t = r_t u_t k_t \tag{A.9}$$

$$(1-\alpha)y_t = w_t n_t \tag{A.10}$$

$$g_t + m_t + r_{t-1}^b b_{t-1} + \tau_t + \delta(u_t)\tau_{kt}k_t = b_t + \tau_{ct}c_t + \tau_{kt}r_t u_t k_t + \tau_{nt}w_t n_t \tag{A.11}$$

$$y_t = c_t + x_t + g_t + m_t \tag{A.12}$$

$$q_t = c_t - \mu c_{t-1}. \tag{A.13}$$

In the above, equations (A.1)–(A.6) are the household's first-order conditions with respect to the decision variables $\{c, x, n, u, k, b\}$, (A.7) is the law of motion for capital, equations (A.8)–(A.10) characterize the firm's profit maximization problem, (A.11) is the government's budget constraint, (A.12) is the aggregate feasibility constraint, and (A.13) adds an auxiliary variable that is used to simplify the household's first-order condition (A.1).

When considering the GHH utility case, we need to make some minor changes to the above system of equations. In particular, equations (A.1), (A.3), and (A.13) change to

$$\lambda_t(1+\tau_{ct}) = d_t^{-\gamma} - \beta\mu E_t d_{t+1}^{-\gamma} \tag{A.14}$$

$$\lambda_t(1-\tau_{nt})w_t = \phi d_t^{-\gamma} n_t^{\chi} \tag{A.15}$$

$$d_t = c_t - \mu c_{t-1} - \phi\frac{n_t^{1+\chi}}{1+\chi}. \tag{A.16}$$

Table 10: Acceptance ratios.

Preferences	Sample	
	1939–54	1955–17
KPR preferences		
Chain 1	0.2528	0.2575
Chain 2	0.2527	0.2576
GHH preferences		
Chain 1	0.2509	0.2583
Chain 2	0.2513	0.2588

A.3 Estimation details

For each preference specification/sample period combination, we run two chains of 2 million replications, using the first 1.75 million as a burn-in. In all cases, we set the tuning parameter to 0.375. Our acceptance rates are in the neighborhood of 0.25, as shown in Table 10.

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